American Journal of Public Health

November 1991, Vol. 81, No. 11 Established 1911

ED KIANKA 86389
CONGRESSIONAL INFORMATION SVC 028
4520 EAST-WEST HIGHWAY
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1015 Fifteenth Street, NW
Washington, DC 20005**

American Public Health Association

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American Journal of Public Health

American Public Health Association

The American Journal of Public Health, published monthly, is the official Journal of the American Public Health Association, 1015 Fifteenth Street, NW, Washington, DC 20005; (202) 789-5600. APHA is a nongovernmental professional society founded in 1872, representing all disciplines and specialties in public health.

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Manuscripts, Correspondence, and Letters to the Editor: Send to Michel A. Ibrahim, MD, PhD, Editor, *American Journal of Public Health,* 1015 Fifteenth St., NW, Washington, DC 20005.

Information on Manuscript Preparation is published in this Journal in January. The Journal publishes reports of original research, demonstrations, evaluations, and other articles covering current aspects of public health. Manuscripts are accepted for consideration with the understanding that they, or their essential substance, have not been published previously or accepted for publication and are not under consideration elsewhere.

Subscription price, individual, is $80/year US, $120/year international; **institutional,** $160/year US, $240/year international. Single copies $10 US, $14 international (surface delivery), $22 international (air delivery). Prepaid only.

Postal Information: Second class postage paid at Washington, DC, and additional mailing offices. Postmaster: Send address changes to American Journal of Public Health, 1015 Fifteenth St., NW, Washington, DC 20005. AJPH allotted $40; TNH allotted $6 of APHA membership dues. Information on individual membership appears elsewhere in this Journal, or contact: Director, Membership Services, at APHA headquarters.

Advertising Sales: Mary Beth Kammann, Advertising Manager, APHA, 1015 Fifteenth St., NW, Washington, DC 20005; (202) 789-5668 or 5600. Advertising is accepted with the understanding that products and services are professionally related and meet the ethical standards of public health practice. Acceptance by the Journal does not indicate or imply endorsement by the Journal of APHA.

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Language of Interview: Relevance for Research of Southwest Hispanics

Bradford Kirkman-Liff, MS, DrPH, and Delfi Mondragón, MHS, MPH, DrPH

ABSTRACT

Background. This paper reports the results of a survey investigating health status, access, satisfaction with care, and barriers to care in Arizona. The major focus is on the association between language of interview and the dependent measures; interviews were conducted in English and Spanish.

Methods. The differences between groups were tested using chi-square statistics for each independent categorical variable; the significance of all the independent variables on each of the dependent variables was tested simultaneously using maximum likelihood logistical regression models.

Results. Language of interview for Hispanic children was a significant variable, more important than ethnicity itself, in determining health status, access, satisfaction with care, and barriers to care; language of interview for Hispanic adults was not a significant measure, but neither was ethnicity. Instead, income affected access to care for adults.

Conclusions. This pattern of results suggests that in the southwestern United States, studies on health status and access to care that use only ethnicity and do not include language of interview may fail to identify populations of Hispanic children who are remarkably more vulnerable. Public health research of Hispanic populations can be more instrumental toward policy improvement if it increases its specificity with this heterogeneous group. Analysis of language of interview has a low cost and a high benefit toward this specification. (*Am J Public Health.* 1991;81:1399–1404)

Introduction

The health of Hispanic Americans has attracted public health research interests for important reasons: poor health status,[1,2] diminished routine use of care,[3-6] and increasing, unprecedented violence.[7,8] The most dominant contextual variables for health, level of income and education, are the lowest in the nation.[9] Moreover, Hispanic population growth exceeds that of all other racial and ethnic groups.[10] In addition to underscoring important empirical issues, methodological inquiries have addressed data uniformity,[11-16] identifier terminology[17-20] and its implications for policy,[21] effects of acculturation,[22,23] and comparative validity of translation.[24-26]

An operationalization for measurement that adds not only to methodological study but to the inferences drawn is language of interview. Among 69 articles published on the topic in the last 10 years, we found none that treated the language of the interview as an analytical variable, even though direct questions about English proficiency and proficiency as a measure of acculturation have been used. Language of interview contributes to specification in research of US Hispanics, particularly in the Southwest, because Hispanics have great genetic, historical, socioeconomic, and political heterogeneity. Furthermore, English language ability sufficient for an interview is often an indirect measure of ability to obtain other health inputs, such as employment, housing, education, nutrition, and health care.

Methods

The data were drawn from a larger survey, commissioned by a major Arizona foundation, investigating health status, barriers to health care, and satisfaction with care in the state. The field work was conducted by Louis Harris and Associates, Inc, using a questionnaire and survey design developed by the investigators, provider and consumer groups, and government agencies in Arizona. The survey was designed to provide reliable information about a representative cross-section of adults and children in all parts of Arizona, excluding those whose places of residence were Indian reservations, military bases, prisons, nursing homes, college dormitories, and mental institutions.

The total survey consisted of 3104 randomly selected adults 17 years of age and over who were interviewed about themselves, and 1113 adults who provided information about a randomly selected child 0 to 16 years of age living in the household. Of the 4217 interviews, 4073 were conducted by telephone during February through May of 1989. In the first stage, interviews were conducted with a representative cross-section of 2000 adults and on behalf of 705 children; in the second stage, additional telephone interviews were conducted with 1004 adults and on behalf of 364 children at or below the poverty level. The child proxy was identified by asking the adult respondent for the

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Bradford Kirkman-Liff is with the School of Health Administration and Policy of Arizona State University. Delfi Mondragón is with the Center for Health Policy and Ethics, Creighton University School of Medicine.

Requests for reprints should be sent to Delfi Mondragón, MHS, MPH, DrPH, Center for Health Policy and Ethics, School of Medicine, Creighton University, 24th and California, Omaha, NE 68178.

This paper was submitted to the journal July 18, 1990, and accepted with revisions May 28, 1991.

identity of the adult in the household who was most familiar with the health care of the child. If this was the adult respondent, then the interview continued; if not, then the interview shifted to that identified person. The oversample was obtained by screening additional households drawn from the same bank of telephone numbers in which low-income persons had been found in the first stage. A household was not interviewed if its income exceeded the poverty level. In both stages of the telephone interviews, rural areas were oversampled to provide better data about less-populated areas, whereas Maricopa County, with 83% of the population, was undersampled.

Lastly, 144 in-person interviews were conducted in households without telephones (100 with adults about themselves and 44 with adults about their children) in order to obtain responses from those populations least likely to have residential telephones—poor, Spanish-speaking Hispanics. These in-person interviews were conducted in selected census tracts in Phoenix, Tucson, and Yuma that had the lowest proportion of households without telephones and the highest proportion of households below the federal poverty line. This resulted in Spanish speakers being more likely to receive a personal interview than English speakers. Hispanic adults accounted for 50% of the in-person adult interviews, compared with 19% of the telephone interviews, before weighting. Hispanic children accounted for 55% of the in-person child interviews, compared with 26% of the telephone interviews, before weighting. This was done to preclude bias from using only telephone interviews, with subsequent underreporting for those without telephones; there are no publicly available data on urban/rural and Hispanic/Anglo telephone penetration rates for simple weighting procedures. We chose instead the documented but smaller potential risks of administration effects of face-to-face[27,28] and different-language interviews.[25] The questionnaire was translated from English to Spanish by consultants in Tucson in order to use appropriate phraseology, and retranslated by other consultants to verify the accuracy of the translation, for both telephone and in-person interviews.

A total of 469 adults and 235 children who labeled themselves Mexican American or Hispanic were identified. They were asked whether they preferred to be interviewed in Spanish or English. Of these, 139 adult interviews and 70 child interviews were conducted in Spanish. All of the tabular analyses reported were performed using weighted observations.

Dependent measures used for recipient health status and access to care were as follows:

- *Current health status:* whether the respondent was in excellent, good, fair, or poor health
- *Disability:* whether the respondent was disabled, handicapped, or had a chronic disease that kept him or her from participating fully in work or other activities
- *Usual source of care:* whether the respondent had one person or place that he or she usually went to when sick or when seeking advice about health
- *Medical care use:* whether, in the 12 months prior to the survey, the respondent had one or more visits to a medical provider and (separately) one or more visits for emergency care
- *Nonreceipt of needed care:* whether the respondent, in the 12 months prior to the survey, thought that he or she needed medical care but did not receive it
- *Refusal of medical care:* whether the respondent, in the 12 months prior to the survey, had been refused health care because of financial or other reasons
- *Access to care:* whether it had become easier or more difficult for the respondent and his or her family to get the medical care they needed
- *Financial problems from illness:* whether illness in the family had been a major financial problem.

The last four items were used only in the adult respondent interviews.

This analysis used the following independent measures:

- *Ethnicity:* Anglo or Hispanic, derived from two separate questions (The term *Anglo* was used as a convenient way of referring to the non-Mexican dominant society, not necessarily indicating British origin. Respondents were asked for their race, from among the categories of White, Black, Asian, Native American, or other. Those who reported "other" were asked for their race: the most common answer was Mexican American, which was recorded. Respondents were separately asked whether they were Hispanic. Individuals who were White and not Hispanic were classified as Anglo, and individuals who identified themselves as Mexican American or White and Hispanic were classified as Hispanic. For the purpose of this analysis, Black and Native American respondents were excluded from all calculations.)
- *Language of interview:* the language in which the interview was completed, recorded separately
- *Employment status of the main wage earner in each household:* whether the main wage earner was employed, unemployed, or not in the labor force (retired, full-time student, homemaker, or disabled)
- *Household income:* classified as poor (below the federal poverty definition), near poor (between federal poverty and 185% of the poverty line), and nonpoor (over 185% of federal poverty)
- *Age of adult respondents:* 17 through 39, 40 through 64, and 65 and older
- *Education of the adult respondent and the main wage earner in each household:* less than a high school education, a high school education, or more than a high school education
- *Residence:* urban (Maricopa or Pima Counties, which include Phoenix and Tucson) or rural Arizona
- *Mode of interview administration:* telephone or in-person

In an exploratory stage, statistical significance of the differences between groups for the reported measures was tested using chi-square statistics for each independent categorical variable. In the final analyses, maximum likelihood logistical regression models were developed that simultaneously tested the significance of all of the independent variables on each of the dependent variables. This process involved forward model construction, to control for possible confounding due to the paucity of independence among variables themselves. In all of these models, the ethnicity and language of interview terms were included to test for their significance, after controlling for all other significant demographic measures. Dummy variables representing different categories within each of the demographic variables were used in the logistical models.

Results

The exploratory cross-tabulations for the nine health status and access measures for adults revealed significant differences between Anglos and Hispanics on seven of these measures, at a 95% confidence level. (These tables can be obtained from the second author.) Six of these measures indicated that Hispanics in Arizona have lower health status and worse access to care than Anglos, before consideration of other demographic variables. These measures included the proportion whose self-described health status was fair or poor, the proportion without a usual source of care, the proportion without an ambulatory care visit in the 12 months prior to the survey, the proportion who sought but were refused care, the proportion whose access to care had become more difficult, and the proportion with financial problems due to illness. The comparison of Hispanics interviewed in English with those interviewed in Spanish revealed that there are significant differences between these two groups on five of these measures. All five of these comparisons indicate that the Hispanics who were interviewed in Spanish have lower health status and worse access to care than Hispanics who were interviewed in English, before consideration of other demographic variables. These measures included the proportion whose self-described health status was fair or poor, the proportion without a usual source of care, the proportion without an ambulatory care visit in the 12 months prior to the survey, the proportion who sought but were refused care, and the proportion with financial problems due to illness.

Similar results were found for the five health status and access measures for children. There are significant differences between Anglos and Hispanics on two of the measures, at a 95% confidence level: the proportion whose proxy-described health status was fair or poor and the proportion who were disabled. Both of these measures indicate that Hispanic children in Arizona have lower health status than Anglos but no worse access to care, before consideration of other demographic variables.

The comparison of Hispanic children whose parents were interviewed in English with those whose parents were interviewed in Spanish indicated that there are significant differences between these two groups on three of these measures. All three of these comparisons indicate that the Hispanic children whose parents were interviewed in Spanish have lower health status and worse access to care than Hispanic children whose parents were interviewed in English, before consideration of other demographic variables. Two of these measures—percentage of children with no usual source of care and percentage of children with no ambulatory care visit in the 12 months prior to the survey—showed no difference between Anglos and

TABLE 1—Maximum Likelihood Logistical Regression Models, Health Status and Access Measures for Adults, Arizona, 1989

Dependent and Independent Variables	β	χ^2	P	R
Health status (0 = excellent, 1 = good, 2 = fair, 3 = poor)				
All Hispanics	.073	0.56	.4549	.000
Spanish-interviewed Hispanics	.279	2.28	.1314	.006
Main wage earner not in labor force	.787	90.37	.0001	.107
Poor household	.296	10.20	.0014	.056
Nonpoor household	−.469	24.20	.0001	−.053
17–39 years old	−.477	37.04	.0001	.033
Less than high school education	.485	26.22	.0001	−.054
More than high school education	−.414	23.73	.0001	−.068
Model $\chi^2 = 647.00$, $P < .0001$, $R = .287$				
Disability (0 = yes, 1 = no)				
All Hispanics	.113	0.57	.4510	.000
Spanish-interviewed Hispanics	−.010	0.00	.9694	.000
Main wage earner not in labor force	−1.492	133.14	.0001	−.210
Main wage earner unemployed	−.502	4.79	.0286	−.031
Nonpoor household	.618	24.67	.0001	.087
17–39 years old	1.207	79.05	.0001	.161
65 years old or older	.648	24.62	.0001	.087
Less than high school education	−.347	8.84	.0029	−.048
Model $\chi^2 = 460.86$, $P < .0001$, $R = .387$				
Usual source of care (0 = yes, 1 = no)				
All Hispanics	−.026	0.05	.8155	.000
Spanish-interviewed Hispanics	.311	2.27	.1317	.008
Main wage earner not in labor force	−.395	16.06	.0001	−.061
Nonpoor household	−.273	10.01	.0016	−.046
17–39 years old	.438	24.51	.0001	.077
Model $\chi^2 = 80.09$, $P < .0001$, $R = .136$				
Had a medical visit (0 = yes, 1 = no)				
All Hispanics	.175	2.61	.1065	.013
Spanish-interviewed Hispanics	.235	1.34	.2466	−.000
Main wage earner not in labor force	−.485	30.88	.0001	−.086
Nonpoor household	−.556	37.44	.0001	−.096
More than high school education	−.241	7.29	.0069	−.037
Rural	.251	9.47	.0021	.044
Model $\chi^2 = 117.03$, $P < .0001$, $R = .165$				
Had an emergency visit (0 = yes, 1 = no)				
All Hispanics	.180	1.27	.1897	.000
Spanish-interviewed Hispanics	.252	0.85	.3553	.000
Main wage earner not in labor force	−.268	7.68	.0056	−.044
Poor household	−.245	6.06	.0138	−.037
Rural	.236	5.86	.0155	.036
Model $\chi^2 = 23.88$, $P = .0002$, $R = .068$				
Needed but did not receive care (0 = yes, 1 = no)				
All Hispanics	.111	0.62	.4297	.000
Spanish-interviewed Hispanics	−.057	0.05	.8202	.000
Poor household	−.308	5.41	.0201	−.036
Nonpoor household	.561	15.58	.0001	.072
65 years old or older	.879	36.26	.0001	.114
Model $\chi^2 = 86.45$, $P < .0001$, $R = .171$				
Refused care (0 = yes, 1 = no)				
All Hispanics	−.087	0.21	.6464	.000
Spanish-interviewed Hispanics	.775	3.31	.0689	.030
Main wage earner unemployed	−.486	4.25	.0393	−.039
Poor household	−1.168	53.23	.0001	−.188
17–39 years old	−.699	19.94	.0001	−.111
Model $\chi^2 = 98.55$, $P < .0001$, $R = .247$				
Access to care (0 = more difficult, 1 = no difference, 2 = easier)				
All Hispanics	−.302	6.47	.0110	−.033
Spanish-interviewed Hispanics	−.043	0.04	.8402	.000
Poor household	−.310	6.86	.0088	−.035
Nonpoor household	.452	12.26	.0005	.050
17–39 years old	−.372	13.39	.0003	−.053
65 years old or older	.587	17.81	.0001	.063
Less than high school education	−.312	8.28	.0040	−.039
Model $\chi^2 = 163.13$, $P < .0001$, $R = .192$				

(Continued)

TABLE 1—Continued

Dependent and Independent Variables	β	χ^2	*P*	*R*
Financial problems from illness (0 = yes, 1 = no)				
All Hispanics	.194	1.70	.1927	.000
Spanish-interviewed Hispanics	−.233	0.92	.3388	.000
Main wage earner not in labor force	−.305	5.47	.0193	−.037
Poor household	−.289	4.93	.0264	−.034
Nonpoor household	.992	38.37	.0001	.121
65 years old or older	.624	15.87	.0001	.074
Less than high school education	−.339	7.68	.0056	−.048
Model $\chi^2 = 153.30$, $P < .0001$, $R = .236$				

TABLE 2—Maximum Likelihood Logistical Regression Models, Health Status and Access Measures for Children, Arizona, 1989

Dependent and Independent Variables	β	χ^2	*P*	R
Health status (0 = excellent, 1 = good, 2 = fair, 3 = poor)				
All Hispanics	.209	1.90	.1678	.000
Spanish-interviewed Hispanics	.584	5.12	.0236	.039
Nonpoor household	−.720	25.22	.0001	−.105
Main wage earner has more than high school education	−.515	13.51	.0002	−.074
Model $\chi^2 = 99.28$, $P < .0001$, $R = .208$				
Disability (0 = yes, 1 = no)				
All Hispanics	−.077	0.06	.8035	.000
Spanish-interviewed Hispanics	−.069	0.02	.8975	.000
Main wage earner not in labor force	−.716	5.06	.0244	−.078
Model $\chi^2 = 4.81$, $P = .1863$, $R = .000$				
Usual source of care (0 = yes, 1 = no)				
All Hispanics	−.383	2.60	.1069	−.025
Spanish-interviewed Hispanics	1.182	12.25	.0005	.105
Main wage earner has less than high school education	.648	10.41	.0013	.096
Model $\chi^2 = 31.95$, $P < .0001$, $R = .168$				
Had a medical visit (0 = yes, 1 = no)				
All Hispanics	−.286	2.47	.1163	−.019
Spanish-interviewed Hispanics	1.069	12.63	.0004	.093
Nonpoor household	−.779	23.96	.0001	−.134
Model $\chi^2 = 43.74$, $P < .0001$, $R = .175$				
Had an emergency visit (0 = yes, 1 = no)				
All Hispanics	.145	0.54	.4619	.000
Spanish-interviewed Hispanics	.572	1.72	.1893	.000
Model $\chi^2 = 3.90$, $P = .6482$, $R = .000$				

all Hispanics but significant differences between those children with English- and Spanish-speaking Hispanic parents. The third measure was the proxy-reported health status, which showed a gradient of declining health status from Anglo children, through Hispanic children with English-speaking parents, to Hispanic children with Spanish-speaking parents.

Analysis of the demographic measures for adults showed significant differences between Anglos and all Hispanics on all six demographic measures and significant differences between Hispanics interviewed in English and those interviewed in Spanish on three measures. There were significant differences between Anglo and all Hispanic children on all four demographic measures and significant differences between Hispanic children whose parents were interviewed in English and those whose parents were interviewed in Spanish on three measures.

The exploratory analysis revealed the need for multivariate statistical techniques that would simultaneously control for the demographic and administration variables in order to assess the Anglo-Hispanic and English-Spanish differences. As the dependent variables were in categorical form, logistical regression was used. Table 1 displays the nine maximum likelihood models for adults. The terms for all Hispanic respondents were significant in only one model: change in access to health care. Hispanic respondents were more likely to report that access to care had become more difficult than other respondents, after controlling for other significant demographic effects. The terms for Spanish-interviewed Hispanic respondents and for mode of administration were insignificant in all models.

Table 2 displays the maximum likelihood models for children. The terms for all Hispanic respondents were insignificant in all five models and significant for Spanish-interviewed Hispanic parents in three models. After controlling for household income and main wage earner education, Hispanic children whose parents were interviewed in Spanish are more likely to be in fair or poor health and less likely to be in good or excellent health than Anglo children or Hispanic children whose parents were interviewed in English. Hispanic children whose parents were interviewed in Spanish were less likely to have a usual source of care than Anglo children or Hispanic children whose parents were interviewed in English. Lastly, Hispanic children whose parents were interviewed in Spanish are less likely to have had a medical visit in the year prior to the survey than Anglo children or Hispanic children whose parents were interviewed in English. The term for mode of administration was insignificant in all models.

In seven of the adult models nonpoor respondents (those with household incomes greater than 185% of the federal poverty definition) had significantly better health status and access to care than the poor and near-poor respondents. In five of the adult models, poor respondents had significantly worse health status and access to care than the near-poor and nonpoor. The poor were also more likely than other respondents to have had one or more emergency room visits. Adults with less than a high school education had significantly worse health status and access to care (as measured by four models) than those who had finished high school or had post-high school education. In two of the children models, their living in a nonpoor household was significantly related to higher health status and access to care.

Discussion

For Hispanics in general, and particularly Hispanics in the southwestern United States, it is important to conduct interviews and analyze separately responses in Spanish for two reasons: to refine translingual and transcultural methodologies and to increase specificity through sampling in this most heterogeneous group. In addition, it presents the opportunity for clinical and social studies of a group with high stress and for study of the long-term effects of political occupation. Each will be discussed.

Comparative validity of translated interview instruments for Spanish speakers in the United States has varied,[22–24] although it appears intact in other, more homogeneous Spanish speakers.[29] Cross-cultural comparisons of health status values have shown greater agreement at the more severe end of the dysfunction continuum and lesser agreement at the less severe end—in the same language.[30] Clearly, internal consistency requires further study. For example, our study would have been greatly enhanced on this dimension if physical examination of respondents had been feasible. At present, however, translated instruments are critical to adequately sample US Hispanics.

Specification of Hispanics in national research is critical because of the significant heterogeneity. There have been waves of immigrants secondary to political change or economic pain from the Caribbean and Central and South America and a steady flow from Mexico. In addition, Puerto Rico as a US territory, has an open-door policy with the United States. Not only does immigration come from diverse places, giving us Black Hispanics and White Hispanics, but the socioeconomic range is great and the political spectrum ranges from extreme right to extreme left. Given this diversity, generalization of Hispanics in national research can clearly be misleading.

Hispanics in the US Southwest have an even more complex history. Briefly, Spain occupied indigent peoples' territory from the Canary Islands in the North Atlantic Ocean, most of the Caribbean Islands, most of Central and South America, and North America—all of Mexico and the territory that is now Texas, California, Utah, Nevada, and parts of New Mexico, Arizona, Colorado, and Wyoming. Mexico overthrew Spain after 300 years, in 1821. In 1848, to end the Mexican War, Mexico agreed to cede the area now the above states to the United States in the Treaty of Guadalupe Hidalgo. The treaty guaranteed Mexicans' land rights, but these were not respected. Many southwestern Hispanics are descendants of area residents from the mid-1500s and earlier, of course—if the indigenous lineage is considered. In the meantime, there have been immigrants from Central and South America and the steady flow from Mexico.

Language of interview focuses on the earliest phase of the Hispano-US trajectory, which all southwestern Hispanics experience either by occupation or by immigration. Spanish monolingualism, in itself, is not a health risk factor but a practical indicator of important risk factors such as diminished education, poverty, and diminished access to care. This is seen as the trajectory progresses. Hispanic immigrants are learning English at the same rate as past generations of other immigrants[31] and are following the classic American pattern of integration into US society.[32] Studies reveal that they accept English as their language in the United States[33] and that the transition from Spanish begins in the immigrant generation. Ninety-five percent of their children speak English, and 50% of this second generation have lost Spanish-speaking ability, speaking English exclusively. These students do not fare much worse than all other students: in 1980, 39% of all Hispanic sophomores were enrolled in remedial courses in English, and 35% of all the nation's sophomores were enrolled in the same remedial classes.[32] While immigrants work at the lowest level jobs, their children move into skilled positions and their grandchildren into professional and managerial positions.[31] Examination of the factors for the 34% wage-offer differential between Mexican-descent men and other men attributes half of this inequality to education and only 2% to fluency in English.[33]

Our study has demonstrated the even greater increased vulnerability in inputs to health and in health care of exclusively Spanish speakers. Four interrelated findings bear elaboration: while reported health status of Spanish speakers and their children is low, the percentage without a usual source of care and without a recent ambulatory visit is significantly lower than the other groups'. We may conclude that the needed care is not obtained. Furthermore, the percentage with financial problems secondary to illness is significantly higher in this group. This, then, forms a vicious cycle: care is not obtained because of financial problems and lack of care causes financial problems. In addition to this hopelessness cycle is the added stress of recent "English only" laws. These are not only confusing, but perceived as menacing by Spanish speakers; often, this perception is accurate as to the roots of the laws. Monolingual Spanish speakers, including those refugees from Central America who have experienced torture and other human rights abuses, can provide insight into specific stressors and stress and into health in general.

The highly stressful early stage of the Hispano-US trajectory and its barriers have historically been overcome, albeit not achieving a status of other US citizens. Perhaps comparability of southwestern Hispanics' health status and other health-related measures is not realistic, given the groups' history. The long-term effects of political occupation, now being debated in other regions of the world, can be studied in southwestern US Hispanics.

Public health research of Hispanic populations can be more instrumental toward policy improvement if it increases its specificity with this heterogeneous group. Analysis of language of interview has a low cost and a high benefit toward this specification. ☐

Acknowledgments

Grateful appreciation is expressed to Delia Saenz, PhD, School of Psychology and Hispanic Research Center, Arizona State University, for comments on a previous draft. This research was supported in part by the Flinn Foundation, Phoenix, Arizona. All opinions and views expressed are solely those of the authors and do not necessarily represent those of the Flinn Foundation.

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Call for Abstracts on Nurse Practitioner Practice

The American Academy of Nurse Practitioners invites abstracts for its 1992 National Conference for Nurse Practitioners to be held in Washington, DC, June 11 to 14, 1992.

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Cancer Incidence in the Puerto Rican–Born Population of Long Island, New York

Anthony P. Polednak, PhD

ABSTRACT

Background. There are apparently no published data on cancer incidence in the Puerto Rican–born populations of the northeastern United States.

Methods. Standardized incidence ratios (SIRs) were calculated for 1980 through 1986 for the Puerto Rican–born population in Long Island (New York).

Results. Significantly reduced SIRs were found for males (SIR = 0.77) but not for females (SIR = 0.91), using expected numbers derived from incidence rates for all areas in the Surveillance, Epidemiology and End Results (SEER) Program (excluding Puerto Rico). Using incidence rates for Puerto Rico to obtain expected numbers, there was evidence for the retention of elevated SIRs for stomach cancer (both sexes) and for significantly elevated SIRs for lung cancer (both sexes), colon-rectum cancer (females), prostate cancer, and breast and uterine corpus cancer. Using rates for SEER areas, the SIRs for lung cancer approached 1.00, in contrast to other US Puerto Rican–born populations.

Conclusion. The data indicate the need for surveys on smoking and other health-related behaviors in the population studied and provide further evidence for heterogeneity in cancer patterns in US Puerto Rican–born populations. (*Am J Public Health.* 1991;81:1405–1407)

Introduction

Migrant studies have contributed significantly to our understanding of changes in cancer rates that reflect environmental influences. Relative to the United States, cancer rates in Puerto Rico are reportedly higher for oral cavity, stomach, and cervix but lower for lung, breast, and corpus uteri.[1] Studies of cancer rates among the Puerto Rican–born population have been limited to mortality for the entire United States[2] and for selected areas such as New York City[3,4] and Illinois.[5] This report presents incidence data on a Puerto Rican–born population in the northeastern United States. According to analysis of data in the 1980 census microdata file,[6] 16.3% of the Puerto Rican–born population of Long Island (i.e., Nassau and Suffolk Counties in New York) was below the poverty level (Table 1); the figure for Puerto Rico was 62.4%.[7] Only 3.2% of the Long Island sample had resided in Puerto Rico in 1975, and only 18.8% did not speak English "well" or "very well" (Table 1).

TABLE 1—Puerto Rican–Born Population of Long Island (New York): 1980 Census

Characteristic	%
Total population	
Males	47.2[a]
Females	52.8[b]
Poverty level (%)	
<100	16.3
100–199	22.9
200+	58.5
Unknown	2.3
Education (age 25+ yrs.)	
High school graduate	43.9
Residence in 1975 (age 25+)[c]	
Same house	49.4
Same state	45.4
Puerto Rico	3.2
Other	2.0
Speaks English (age 25+)	
Only	9.3
Very well	43.7
Well	28.2
Not well	18.8

Note. Data calculated from 1980 census 5% microsample files.[6]
[a] n = 9440.
[b] n = 10 500.
[c] Based on a subsample.[6]

Methods

Data were obtained on all incident cancer cases (without personal identifiers) diagnosed among Puerto Rican–born residents of Long Island during 1980 through 1986 and reported to the population-based New York State Cancer Registry; 1986 was the latest year for which data were available at the time analyses were conducted. Cancers had been coded to the 9th Revision of the International Classification of Diseases.

Estimated numbers of Puerto Rican–born persons in the Nassau–Suffolk standard metropolitan statistical area[6] in 5-year age groups for each sex were multiplied by age- and sex-specific cancer incidence rates for all areas of the Surveillance, Epidemiology and End Results

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The author is with the Department of Epidemiology and Public Health, Yale School of Medicine. Part of this work was done while the author was with the Department of Preventive Medicine, School of Medicine, State University of New York at Stony Brook.

Requests for reprints should be sent to Anthony P. Polednak, PhD, Department of Epidemiology and Public Health, Yale School of Medicine, PO Box 3333, 60 College Street, New Haven, CT 06510.

This paper was submitted to the journal December 26, 1990, and accepted with revisions May 7, 1991.

TABLE 2—SIRs for Cancer in the Puerto Rican–Born Population of Long Island (New York) in 1980–1986, Using Rates from the SEER Program (excluding Puerto Rico) and from Puerto Rico: Males

		Using SEER Rates		Using Puerto Rico Rates	
Site	Observed No.	Expected No.	SIR (95% CI)	Expected No.	SIR (95% CI)
All sites	177	228.5	0.77 (0.66, 0.89)*	134.8	1.31 (1.13. 1.52)*
Oral	13	11.4	1.14 (0.61, 1.95)	15.9	0.82 (0.44, 1.40)
Stomach	12	6.7	1.79 (0.93, 3.13)	11.8	1.01 (0.52, 1.77)
Colon–rectum	16	31.1	0.51 (0.29, 0.83)*	12.7	1.26 (0.72, 2.04)
Lung	39	46.6	0.84 (0.60, 1.14)	12.9	3.02 (2.16, 4.11)*
Prostate	31	35.5	0.87 (0.59, 1.24)	18.7	1.66 (1.13, 2.37)*
Bladder	7	15.2	0.46 (0.18, 0.95)*	6.7	1.05 (0.42, 2.16)

Note. SIRs = standardized incidence ratios; SEER = Surveillance, Epidemiology and End Results. Data are shown only if the observed number, or the expected number (using SEER rates), was 10 or more. Adjustment was not made for estimated Puerto Rican-born cases among all cases with unknown place of birth in the Registry; such adjustment would increase slightly the SIRs reported in the tables, but use of 1980 Census data to estimate the population for 1980-86 has undoubtedly already resulted in some overestimation of the SIRs.

*$P < .05$.

TABLE 3—SIRs for Cancer in the Puerto Rican–Born Population of Long Island (New York) in 1980–1986, Using Rates from the SEER Program (excluding Puerto Rico) and from Puerto Rico: Females

		Using SEER Rates		Using Puerto Rico Rates	
Site	Observed No.	Expected No.	SIR (95% CI)	Expected No.	SIR (95% CI)
All sites	218	239.3	0.91 (0.79, 1.05)	136.0	1.60 (1.39, 1.84)*
Stomach	9	3.4	2.64 (1.21, 2.30)*	7.1	1.26 (0.58, 2.39)
Colon–rectum	29	26.7	1.09 (0.73, 1.57)	12.6	2.30 (1.54, 3.31)*
Lung	23	23.5	0.98 (0.62, 1.47)	5.4	4.25 (2.69, 6.38)*
Breast	62	79.0	0.78 (0.60, 1.01)	36.8	1.69 (1.31, 2.20)*
Cervix	13	9.5	1.37 (0.73, 2.34)	16.6	0.78 (0.41, 1.33)
Corpus	18	16.9	1.07 (0.63, 1.69)	7.9	2.29 (1.36, 3.62)*
Ovary	8	10.7	0.75 (0.32, 1.48)	5.5	1.45 (0.62, 2.86)

Note. SIRs = standardized incidence ratios; SEER = Surveillance, Epidemiology and End Results. Stomach cancer was included for comparison with data for males.

(SEER) Program of the National Cancer Institute, excluding Puerto Rico, for 1981 through 1985.[8] Expected numbers were also obtained by using age-, sex-, and site-specific cancer incidence rates for Puerto Rico (available for 1978 through 1982).[9] Age-adjusted incidence rates for Puerto Rico changed little between 1980 through 1984 and 1985 through 1987.[10] Statistical analysis involved calculation of 95% confidence intervals (CIs)[11] on standardized incidence ratios (SIRs).

Results

For males (Table 2) the SIR for all sites was significantly lower than 1.00 on the basis of SEER rates, but significantly greater than 1.00 on the basis of rates for Puerto Rico. The SIR for stomach cancer (using SEER rates) was 1.79, with no evidence for a change from numbers expected using rates for Puerto Rico. Using SEER rates, significantly reduced SIRs were evident for colon–rectum and bladder cancer. SIRs for prostate and lung cancer were slightly less than 1.00 using SEER rates but were significantly elevated on the basis of cancer rates for Puerto Rico.

For females (Table 3) the SIR for all sites was only slightly and insignificantly reduced using SEER rates, and significantly elevated on the basis of rates for Puerto Rico. The SIR for stomach cancer was even higher than that for males. In contrast to males, the SIR for colon–rectum cancer was greater than 1.00 using SEER rates and significantly elevated on the basis of rates for Puerto Rico. The SIR for lung cancer was nearly 1.00 (using SEER rates) and significantly elevated on the basis of rates for Puerto Rico. The SIR for breast cancer was slightly less than 1.00 using SEER rates and, along with uterine corpus cancer, significantly elevated on the basis of rates for Puerto Rico. The SIR for cervix cancer was slightly greater than 1.00 using SEER rates and less than 1.00 on the basis of rates for Puerto Rico.

Discussion

Potential inaccuracies in the SIRs are presumably due mainly to errors in estimating the population at risk and hence the expected numbers of cancers. Census data from 1980 should be more accurate for the 1980 through 1982 period than for 1980 through 1986, and observed cancers for 1980 through 1982 were compared with expected numbers based on incidence rates for SEER (excluding Puerto Rico) for 1978 through 1981.[12] For all cancers the SIR for 1980 through 1982 was significantly reduced for males (i.e., 0.72; 95% CI = .56, .92) but not for females (i.e., 0.80; 95% CI = .64, 1.00). SIRs for 1980 through 1982 based on at least 10 observed cases were 0.76 (15 observed vs 19.85 expected) for lung cancer and 0.92 (13 observed vs 14.16 expected) for prostate cancer in males, and 0.62 (19 observed vs 30.84 expected) for breast cancer and 1.22 (14 observed vs 11.45 expected) for colon–

rectum cancer in females, or similar to the SIRs in Tables 2 and 3.

For 1980 through 1986, males but not females had significantly reduced SIRs, on the basis of SEER cancer rates, for all cancers and for several sites. For females, both lung and reproductive system sites, with the exception of cervix, showed high SIRs calculated by using cancer rates for Puerto Rico. The findings are clearly consistent with other studies of migrants,[13] showing retention of high risks of stomach cancer (characteristic of the homeland) and a transition toward higher risks in the United States for certain other cancers such as those of the lung, prostate, and female reproductive organs.

The finding of an only slightly elevated SIR for cervical cancer (using SEER rates) may reflect the higher socioeconomic status (SES) of the Puerto Rican–born population in Long Island vs other areas where relatively high mortality for cervical cancer has been reported for Puerto Rican immigrants.[3–5] Selective migration, related to such characteristics as SES, is possible but difficult to document. The higher SES of the Puerto Rican–born population in Long Island (vs Puerto Rico) is most likely due to improvement in SES after migration from Puerto Rico, with subsequent movement to Long Island (mainly from New York City). Increasing SES may result in changes in environmental factors such as dietary and sexual–reproductive patterns, leading to increased risks of certain cancers (e.g., uterine corpus and breast) and a reduced risk of cervical cancer; the explanation of patterns for lung cancer may be more problematic.

Changes in smoking behavior among migrants are difficult to assess because data on smoking prevalence in Puerto Rico are limited to a 1965 through 1968 survey of men 45 through 64 years old,[14] and age at migration was unknown for the present group of migrants. In the 1982 through 1984 Hispanic Health and Nutrition Survey, Puerto Rican men and women had smoking rates higher than those of non-Hispanic Whites, but data have not been reported for the Puerto Rican–born subgroup.[14,15] The effect of increasing SES and/or acculturation on smoking habits and lung cancer rates in specific Hispanic subpopulations[14] requires further study, in view of the present findings on lung cancer in a higher income group of Puerto Rican–born US residents. ☐

Acknowledgments

This research was supported in part by Grant 5 P01 CA42101 from the National Cancer Institute to the Cancer Prevention Research Unit at the Yale School of Medicine. Part of these data were provided by the Cancer Control Bureau of the New York State Department of Health; the author assumes full responsibility for the analysis and interpretation of the data.

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Injuries in Working Populations: Black-White Differences

Diane K. Wagener, PhD, and Deborah W. Winn, PhD

ABSTRACT

Background. Although "accidents and adverse effects" mortality is higher among Blacks than Whites, annual injury rates reported in the National Health Interview Survey (NHIS) are lower among Blacks. We evaluated the influence of sociodemographic risk factors on injury rates among working adults.

Methods. NHIS data from 1983 through 1987 for currently working adults were used. Methods were developed to estimate standard errors using data from different sample frames and sample sizes.

Results. Working Blacks had fewer reported injuries requiring medical attention or restriction of usual activities than working Whites (22.0 vs 27.0 per 100 persons per year). The difference was pronounced among younger adults in both sexes and among both poor and nonpoor. However, age, sex, and income could not completely explain racial differentials. "At-work" injury rates (36% of all injury episodes) were similar for Blacks and Whites (9.2 vs 9.9 per 100 persons per year), except low-income Blacks and Blacks in service or blue-collar occupations had nonsignificantly smaller at-work injury rates.

Conclusions. Possible reporting biases could not be completely eliminated. However, available evidence does not rule out a true difference in injury rates by race, highlighting the complexity of understanding the etiology of injuries and, hence, developing public health programs to prevent injuries.(*Am J Public Health.* 1991;81:1418–1414)

Introduction

Injuries are a major cause of death among young adults and adults in their middle years.[1] More than one third of deaths among adults aged 20 to 34 years are due to accidents and adverse effects.[2] In contrast, 12% of deaths to adults aged 35 to 49 years and only 3% of deaths to adults aged 50 to 64 years are due to accidents and adverse effects. In addition to these deaths, each year there are 33.8 million nonfatal injuries to adults requiring medical attention or restriction in activity.[3,4] This report focuses on episodes of nonfatal injury.

In general, working adults have lower injury rates than other segments of the population.[3] However, injuries do have significant consequences for the working population. Injuries cause the loss of more working years of life than all forms of cancer and heart disease combined.[1] In 1980, injuries and poisonings were the leading category for medical costs among working age individuals in the United States.[5] Further, injury rates among workers are rising. From 1983 to 1989, the work-related injury rates reported by the Bureau of Labor Statistics' Annual Survey of Occupational Injuries and Illnesses steadily increased from 7.5 per 100 full-time workers to 8.2.[6,7] Likewise, yearly work loss due to injuries has increased (from 57.2 days per 100 workers to 74.2). These data are based on reports by employers that are subject to underreporting.[8]

Injuries occur in some population subgroups more than others.[9] Whites report more injuries than Blacks[10] but are far less likely to die of injuries than Blacks.[11–13] Injury rates are higher at younger ages.[3,4] To measure morbidity in working populations, regardless of its association with the workplace or employer-based reporting systems, and to examine the patterns of injuries in workers related to known important parameters such as race, we examined the incidence of injuries among US workers using data from the National Health Interview Survey (NHIS).

Methods

The NHIS is a continuing nationwide sample survey using personal household interviews conducted by the National Center for Health Statistics of the Centers for Disease Control.[4] Questionnaire information is collected on demographic and personal characteristics, illnesses, injuries, impairments, chronic conditions, health resource use, and other current health topics. The selection of households is based on a multistage probability sampling plan permitting national estimates for the noninstitutionalized civilian population of the United States. Over the years covered in this report, the response rate for the survey was between 95% and 98%.[4]

For this report, data from the survey years 1983 through 1987 were combined. Only currently employed adults 18 through 64 years of age were included. (Currently employed adults 65 years and over accounted for only 6952 out of a total of 216 528 interviews of all working

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Diane K. Wagener is with the Division of Epidemiology and Health Promotion, and Deborah W. Winn is with the Division of Health Interview Statistics, National Center for Health Statistics, Centers for Disease Control, Atlanta.

Requests for reprints should be sent to Diane K. Wagener, PhD, National Center for Health Statistics, Room 1070, 6525 Belcrest Road, Hyattsville, MD 20782.

This paper was submitted to the journal July 2, 1990, and accepted with revisions April 2, 1991.

adults.) "Currently employed" included persons reporting that at any time during the 2-week period covered by the interview they either worked at or had a job or business. This included both paid and unpaid work. Those who were temporarily absent from a job or business were considered as currently employed if they expected to work as soon as the particular event causing the absence (e.g., vacation, strike, or illness) no longer existed.

An injury or injury condition was any condition classified according to the nature-of-injury codes 800 through 999 in the *International Classification of Diseases*, 9th version (ICD-9), modified for the NHIS.[14] An "episode of injury" was any accidental or nonaccidental violence causing an injury requiring medical attention or at least a half day of restricted activity, including bed days, work-loss days, or cut-down days. The recall period was 2 weeks prior to the interview. A person may have had more than one episode of injury during the recall period. Further, more than one type of injury could result from a single episode of injury. For the purposes of this report, "injury" will refer to injury episodes. The number of injury episodes during the 2-week recall period was multiplied by 26. Therefore, "injury rate" refers to the annual rate of episodes of injury.

TABLE 1—Estimated Number of Currently Employed Adults, Aged 18–64 Years, by Several Demographic Characteristics: National Health Interview Survey, 1983–1987

Demographic Group	Estimated Population Size, in Thousands		No. of Interviews
	No.	%	
Total	103 622		209 576
Age (in years)			
18–44	74 963	72.3	149 513
45–64	28 659	27.7	60 063
Men	57 715	55.7	114 617
Women	45 907	44.3	94 959
Whites	90 312	87.2	180 326
Blacks	10 422	10.1	23 704
Other	2388	2.7	5546
Family income			
< $10 000	8379	8.1	17 128
$10 000–$19 999	18 934	18.3	38 926
≥ $20 000	69 156	66.7	138 924
Missing	7153	6.9	4600
Occupation[a]			
White-collar	57 690	55.7	115 936
Blue-collar	28 723	27.7	57 974
Farmer or farm worker	2933	2.8	6046
Service	12 458	12.0	25 813
Missing	1818	1.8	3805
Education			
< 12 years	57 981	60.0	118 470
Some college	35 076	33.8	69 919
Some postgraduate	9548	9.2	19 098
Missing	1017	1.0	2089

[a] Occupation classes include the following: White-collar: executive, administrative, and managerial; professional specialty; technicians and related support; sales; administrative support, including clerical. Blue-collar: precision production, craft and repair; machine operators, assemblers and inspectors; transportation and material moving; handlers, equipment cleaners, helpers, and laborers (except farm). Farm workers: farming, forestry, and fishing. Services: private household; protective service; service occupations, except protective and household.

For this report, data from 5 survey years were combined to enhance the stability of estimates for demographic subgroups. In 1985 the sampling frame for the NHIS was changed to include, among other features, oversampling in geographic areas with a high density of Blacks. Further, because of lack of funds, the numbers of households in the sample had to be reduced by one quarter in 1985 and by one half in 1986.

The NHIS is a clustered multistage probability sample; therefore, the Statistical Analysis System (SAS) procedure SESUDAAN[15] was used to compute rates, means, and their standard errors and to compute the effect of the complex survey design on the standard error, known as the "design effect".[16] Average annual rates for the 5-year period were estimated by pooling episodes and persons from all 5 years and, hence, both survey designs. The variance was, therefore, the weighted sum of the variance due to the 1983 through 1984 survey design period and the variance due to the latter 1985 through 1987 period. The variance for 1983 through 1984 could be calculated by using SESUDAAN. However, because sample sizes changed during 1985 through 1987, the variance for 1985 through 1987 could not be calculated in the usual manner and so was estimated using two approaches. The first approach assumed that the design effect was the same for both time periods and equal to the design effect for the sample collected in 1983 through 1984, deff_o. The variances for the 1985 through 1987 period were then estimated by multiplying the simple random sampling variance by deff_o. The second approach assumed that the coefficient of variation was constant over both time periods. Therefore, using the estimates for the rates in each time period and the variance for the first 2 years, the variance for the last 3 years could be estimated.

The design effects for injury-related parameters by demographic subgroups were generally between 0.7 and 1.4, which suggests only a moderate influence of the complex design on the standard errors. Standard errors presented here use the second—coefficient of variation—approach because (1) there was less than a 10% difference in the magnitudes of standard errors yielded by the two procedures, and (2) the coefficient of variation approach tended to yield larger (and hence more conservative) standard error estimates, especially for smaller subpopulations.

Results

For the following analyses, the population included currently employed adults, aged 18 through 64 years. Table 1 gives the estimated population sizes and numbers of people interviewed by several demographic characteristics. Among the Black population, slightly more working adults were in the younger group (75%) or were women (50%).

Although the number of interviews was quite large (209 576), only 2016 working individuals reported an injury requiring medical attention or restricted activity in the previous 2 weeks, an annual weighted estimate of 27 382 000 injuries in the US working population. Table 2 gives the weighted numbers of injuries and rates

TABLE 2—Annual Number and Rate of Episodes of Injury among Currently Employed Persons, Aged 18–64 Years, by Selected Demographic Characteristics: National Health Interview Survey, 1983–1987

Demographic Group	Number of Episodes of Persons Injured (in thousands)	Rate (per 100 persons/yr)	Standard Error (per 100 persons/yr)
Total	27 382	26.42	0.47
Age (in years)			
18–44	22 128	29.52	0.61
45–64	5254	18.33	0.61
Men	16 708	28.95	0.70
Women	10 674	23.25	0.62
Whites	24 569	27.20	0.49
Blacks	2287	21.95	1.01
Other	526	18.20	0.82
Income			
< $10 000	3427	40.90	2.41
$10 000–$19 999	5736	30.30	1.04
≥ $20 000	16 858	24.38	0.48
Occupation[a]			
White-collar	12 433	21.62	0.50
Blue-collar	10 398	36.20	1.11
Farmer or farm worker	961	32.75	3.80
Service	3210	25.58	1.35
Education			
≤ 12 years	16 070	27.72	0.61
1–4 years of college	9044	25.89	0.75
Postgraduate	2039	21.84	1.20

[a] Occupation classes include the following: White-collar: executive, administrative, and managerial; professional specialty; technicians and related support; sales; administrative support, including clerical. Blue-collar: precision production, craft and repair; machine operators, assemblers, and inspectors; transportation and material moving: handlers, equipment cleaners, helpers, and laborers (except farm). Farm workers: farming, forestry, and fishing. Services: private household; protective service; service occupations, except protective and household.

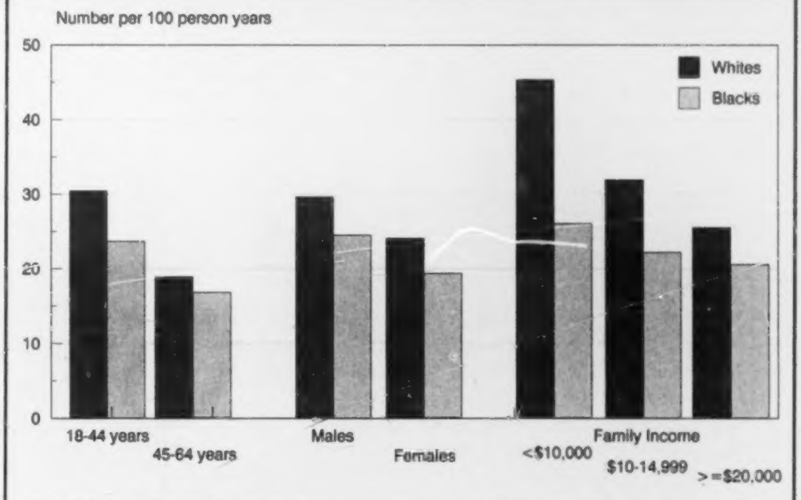


FIGURE 1—Injury rates among currently employed adults; race by age, sex, and income. (Source: National Center for Health Statistics, National Health Interview Survey, 1983–1987)

per 100 persons per year for the various demographic subpopulations listed in Table 1. For all characteristics, the rates were significantly different between pairs of subpopulations. (For race, "Whites" was used as the reference population; for income, "≥ $20,000"; for occupation, "white-collar"; and for education, "postgraduate.") Overall, approximately one in four working adults each year had an injury requiring medical attention or restricted activity. Approximately 89% of all injuries were medically attended (89% of injuries among Whites and 87% among Blacks), and 55% were associated with at least a half day of restricted activity (55% among Whites and 61% among Blacks). Injury rates were higher among the younger male, lower income, and White populations. Blue-collar and farm-related occupations had higher injury rates.

Figure 1 shows the injury rates for age, gender, and family income groups separately for Whites and for Blacks. In all comparisons except workers aged 45 through 64 years, working Blacks had fewer injuries per person per year.

Within age groups, differences between injury rates also were observed by gender. Among workers aged 45 through 64 years, there was no difference between men and women (18.7 per 100 persons per year vs 17.8), whereas, in the younger age group 18 to 44 years, men had a significantly higher reported injury rate (33.0) than women (25.3). When racial groups were compared separately for men and for women by age group, Black men in the younger age group had lower injury rates than White men (26.5 vs 33.9), but there was little difference by race among older men (18.8 vs 19.1). Among women, Blacks had lower rates in both age groups (20.9 vs 26.1 among women aged 18 to 44 years and 14.8 vs 18.4 among women aged 45 through 64 years).

Because the distribution of jobs differs between Whites and Blacks, Table 3 shows the injury rates by four major occupational classes and by race. There was no difference between Whites and Blacks among white-collar workers. There were too few Black workers in farm or farm-related occupations to reliably estimate the annual injury rate. Blacks had significantly lower injury rates among blue-collar workers and service workers. Within these occupational classes, there were racial differences in the types of jobs. Among service workers, 5% of Whites and 10% of Blacks worked in private households. Among the blue-collar workers, 12% of Whites and 19% of Blacks were handlers or equipment cleaners.

To determine whether the overall higher injury rates among blue-collar workers and service workers were found disproportionately among individuals with lower family incomes, the injury rates were further classified by family income (see Table 3). Low-income Blacks were more often working in service occupations (30%) than low-income Whites

(18%). There was little racial difference in the proportions employed in blue-collar occupations (33% of low-income Blacks and 36% of Whites). The injury rates among Blacks were smaller than among Whites for all occupation–income groups, except white-collar workers.

In the NHIS, injuries are grouped in four general classes: moving-motor-vehicles injuries, with traffic accidents as a subclass; accidents occurring while at work; accidents occurring in or adjacent to the home; and other accidents. These classes are not mutually exclusive. In the currently employed adult population, the distribution among classes of accidents was 12%, 36%, 25%, and 34%, respectively. Approximately 13% of the moving-motor-vehicle accidents occurred at work.

Because "at-work" injuries were the largest class of injuries and because all persons in this study population were currently employed, the at-work injury rates were calculated for the demographic subgroups (Table 4). In general, the at-work injury rate pattern was similar to all-injury rates. That is, rates were higher among the younger, male, and lower-income groups. However, in contrast to the all-injury rates, there was little difference between racial groups in at-work injury rates. Figure 2 shows these rates by racial groups for age, sex, and family income separately. Only among the low-income population was the Black at-work injury rate appreciably, though not significantly, lower than the White at-work injury rate. Approximately 93% of the at-work injuries were medically attended (93% among Whites and 94% among Blacks), and 54% were associated with at least a half day of restricted activity (54% among Whites and 59% among Blacks).

The rates for at-work injuries were calculated for the four major occupational groups by race (Figure 3). None of the comparisons by race were statistically significant. In general, the differences in rates of at-work injuries between racial groups were smaller than racial differences in rates of all injuries.

TABLE 3—Injury Rates in Major Occupation[a] Classes, Ages 18–64 years, by Race and Family Income: National Health Interview Survey, 1983–1987

Income and Race	White-Collar	Blue-Collar	Farm	Services
All income levels				
All races	21.82	36.20	32.75	25.58
Whites	21.84	37.95	30.24	28.27
Blacks	21.70	25.57	64.06[b]	14.77
Less than $20 000				
All races	26.60	39.83	39.22[b]	28.72
Whites	27.19	43.38	36.25[b]	31.82
Blacks	29.01	20.64	75.66[b]	19.19
At least $20 000				
All races	21.31	34.93	23.77[b]	19.13
Whites	21.62	36.05	23.86[b]	19.24
Blacks	23.47	27.64	35.97[b]	13.22

[a]Occupation classes include the following: White-collar: executive, administrative, and managerial; professional specialty; technicians and related support; sales; administrative support, including clerical. Blue-collar: precision production, craft and repair; machine operators, assemblers, and inspectors; transportation and material moving; handlers, equipment cleaners, helpers, and laborers (except farm). Farm workers: farming, forestry, and fishing. Services: private household; protective service; service occupations, except protective and household.

[b]Relative standard error greater than 30%.

TABLE 4—Annual Number and Rate of Episodes of Injury at Work among Currently Employed Persons, Aged 18–64, by Selected Demographic Ch[illegible]istics: National Health Interview Survey, 1983–1987

Demographic Group	Number of Episodes of Persons Injured (in thousands)	Rate (per 100 persons/yr)	Standard Error (per 100 persons/yr)
Total	9967	9.62	0.29
Age (in years)			
18–44	8114	10.82	0.36
45–64	1852	6.46	0.37
Men	7335	12.71	0.44
Women	2631	5.73	0.37
Whites	8897	9.85	0.31
Blacks	959	9.20	0.86
Other	112	3.87	
Income			
< $10 000	1236	14.75	1.42
$10 000–$19 999	2587	13.66	0.75
≥ $20 000	5664	8.20	0.24
Occupation[a]			
White-collar	2759	4.78	0.26
Blue-collar	5494	19.13	0.78
Farmer or farm worker	481	16.40	2.19
Service	1177	9.45	1.03
Education			
≤ 12 years	6873	11.85	0.42
1–4 years of college	2508	7.15	0.40
Postgraduate	503	5.26	0.68

[a]Occupation classes include the following: White-collar: executive, administrative, and managerial; professional specialty; technicians and related support; sales; administrative support, including clerical. Blue-collar: precision production, craft and repair; machine operators, assemblers, and inspectors; transportation and material moving; handlers, equipment cleaners, helpers, and laborers (except farm). Farm workers: farming, forestry, and fishing. Services: private household; protective service; service occupations, except protective and household.

Discussion

This study confirmed some previous findings in the injury epidemiology literature, but also generated some intriguing results. Injuries to US working adults are common, affecting a quarter of US workers every year. Reported injuries occur more frequently among the poor, among young men, and among the White population. Total reported-injury rates for working Whites exceeded those for working Blacks (27 vs 22 per 100 persons per year, respectively), but injuries occurring at work were not different between races (9.9 and 9.2, respectively).

In contrast to these findings for nonfatal injuries, injury death rates are about 18% higher in Blacks for injuries occurring at work[12,13] and for many specific causes of injury death regardless of whether the injury occurred at work.[11] Specific types

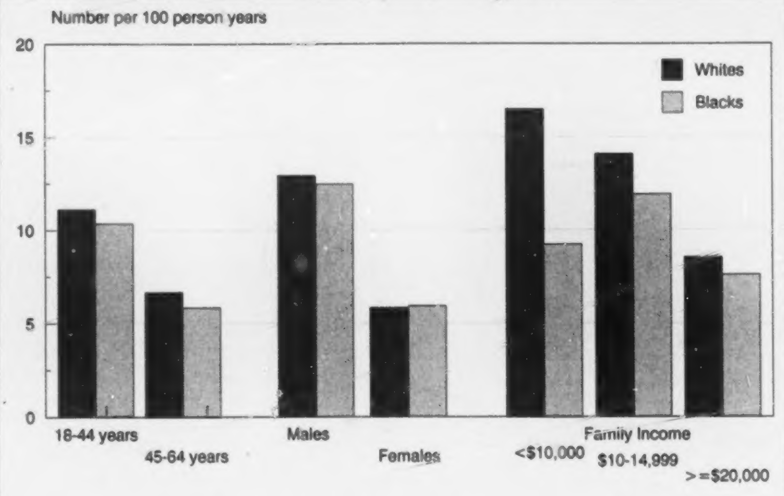
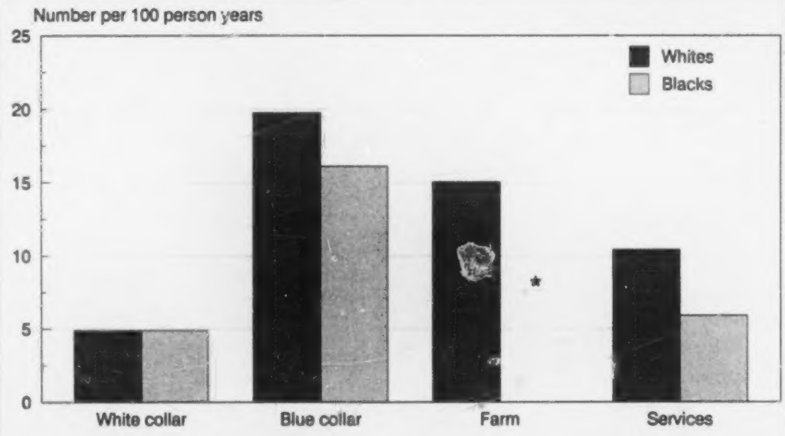


FIGURE 2—At-work injury rates among currently employed adults; race by age, sex, and income. (Source: National Center for Health Statistics, National Health Interview Survey, 1983–1987)



* not shown because relative standard error greater than 30 percent.

FIGURE 3—At-work injury rates among currently employed adults, by occupation and race. (Source: National Center for Health Statistics, National Health Interview Survey, 1983–1987)

of injuries (including nonfatal injuries) show different racial patterns. Hospital- and clinic-based studies have reported more Whites among motor vehicle injury or spinal cord injury victims,[17,18] and more non-Whites among penetrating injury or ocular trauma victims.[17,19] The demographic distributions of these patients, however, may reflect the populations served and not necessarily the rates of injuries in the population subgroups.

Major national surveillance systems that include nonfatal injury information, such as the Bureau of Labor Statistics' Annual Survey of Occupational Injuries and Illnesses and the Consumer Product Safety Commission's National Electronic Injury Surveillance System, either do not collect race or ethnicity data or do not routinely tabulate data by race. The NHIS is a source of national nonfatal injury information that provides both racial and ethnic information. This report and an analysis of data from the National Health Surveys from June 1959 through June 1961[9] (predecessors to the NHIS) show higher nonfatal injury rates for Whites, except for injuries occurring at work. In both analyses, the reported injury rates for Whites were about 23% higher than for Blacks (in this report) or for non-Whites (in the earlier report), although the rates have doubled over the time period. Reported injuries occurring at work have also doubled, but the racial difference has remained only about 5%. The authors of the earlier report hypothesized that racial differences in all-injury rates could be due to socioeconomic status, "culture" (including patterns of use of medical facilities), and nature of the work performed.[9] Some of these factors are evaluated in the present report.

Young (ages 18 to 44), working White men had the highest reported injury rates, significantly higher than their Black counterparts. Injury rates were highest among blue-collar workers. Among blue-collar and service workers, White rates exceeded those for Blacks, and this remained true when these occupational groups were further subdivided into income categories. Injuries occurring at work account for more than one third of all injuries occurring to working persons. The pattern of at-work injuries showed little of the racial differences observed for injuries in general, with the exception that, among the poor, at-work injury rates for Whites exceed those for Blacks.

In the NHIS, an event cannot be classified as an injury episode unless medical treatment was sought or there was at least one half day of restricted activity. The higher reported rate of injuries among Whites could be due to a truly higher rate, better access to medical care for Whites compared with Blacks, a greater opportunity for Whites to cut down on routine activities because of illness, or a higher probability that Whites will report an injury episode in this interview survey format.

A conclusion that Whites have higher rates of injuries than Blacks would be strengthened if it could be shown that other factors associated with race and the occurrence of injuries did not account for the findings. All of the factors we examined that were associated with greater injury reporting—lower income, lower education, and young age—were inversely associated with being White, suggesting that confounding could not account for the results.

Assessing access to medical care is complex. Other data from the NHIS[10] suggest that reported rates for Whites exceed those for Blacks for some health indicators that reflect contact with the medical care system: the incidence of acute conditions (defined in part by contact with a medical care provider) and physician contacts per year. In this report, the percentage of episodes associated with medical care visits did not differ between Whites and Blacks, providing some indication that access to care was not responsible for our findings.

Family income[20] and health insurance coverage[21,22] are two important indicators of access to medical care. Blacks are more likely than Whites to be uninsured, potentially limiting access to medical care.[21,22] Insurance coverage data were unavailable for most of the survey years covered in this report, but income data were collected in all 5 survey years. Individuals with low family incomes might be expected to have less access to medical care and be less able to take time off from their activities, such as work. Therefore, it might be hypothesized that injury reporting would be lower in the poorer group even if their true rates were equal to or higher than those in other income groups. However, workers with family incomes of less than $20 000 had higher reported injury rates than higher income workers, and those with family incomes under $10 000 had injury rates higher still. Even among the poorest—where, regardless of race, income might severely influence taking time off or getting medical care—White rates exceeded rates for Blacks. Moreover, 88% of the injuries to low-income individuals were medically attended, nearly equal to the 90% rate for higher income individuals. Equal proportions of doctor visits in both income groups, however, does not necessarily indicate similarities in access to care. Data were not available to determine whether injuries considered severe enough to seek medical attention by one group were medically attended by individuals in another group.

Racial differences in the opportunities to cut down on routine activities or to stay in bed could explain the higher reported injury rates among Whites. Blacks may be more likely to have jobs in which staying away from work for reasons of illness carries greater financial consequences, although family income did not account for the racial differences reported here. Alternatively, family and nonwork responsibilities may make routine activity restriction difficult. A greater percentage of Blacks are in single-parent households[23] where opportunities for restricting activity may be limited.

If Whites have a higher probability of reporting health problems in an interview format, then this may account for the findings. Greater reporting by Whites probably would lead to Blacks having reported more severe conditions on average, but this could not be evaluated with these data. The National Center for Health Statistics is conducting a study of the reliability of reporting of health conditions using the NHIS interview and medical record data that may help evaluate racial differences in health event reporting.

Evaluation of potential biases is difficult. However, from the discussion above, it appears that the higher rate of injuries among Whites cannot be explained by sociodemographic correlates of race. Available indicators of access to care (e.g., percentage of injuries medically attended) and analyses by income level suggest that availability of health care did not account for the findings. No data were available to address whether racial differences in opportunities for activity restriction or willingness to report injuries had an impact on these results.

This study highlights the importance of analyzing separately injury patterns by the setting of the injury episode because the incidence of injuries at work did not show the same large differences by race for total injuries. Perhaps differences in access to care or opportunities for cutting down on routine activities generally are "leveled" by the workplace. Individuals working in the service occupations (e.g., personal service workers) are less likely to be provided health insurance packages[22] or may be offered less workplace health care, sick leave, or work-hour flexibility. This may explain the lower at-work injury rates among service workers compared with blue-collar workers. Further, our data show that Blacks are more likely to be in these occupations.

Despite their limitations, interviews with the victims of injuries would seem to offer the best opportunity for studying racial differences in nonfatal injury occurrence. These data are less dependent on access to care compared with reporting systems based on emergency room, compensation data, or physician records. However, this focused examination of racial differences in injury reporting suggests that methodological research be conducted on approaches to collect injury data via questionnaires. It will be important to reexamine existing definitions of injuries used in interview surveys to determine the extent to which these capture the injuries of greatest public health importance. Improved assessments of access to care and of barriers to reporting are needed in epidemiologic studies. Measuring severity of injuries should provide better insight into the public health aspects of injury and disability.

Injury epidemiology is an important public health concern, as evidenced by its inclusion as an emphasis area in the national health objectives for the years 1990 and 2000.[7,24] The workplace is an effective arena for health promotion programs. Therefore, understanding injury patterns within sociodemographic subgroups of the US working population is important. However, it is also important to have better information on injury patterns within sociodemographic subgroups of the entire US population. ☐

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Health and Population Data for Developing Countries Available through USAID Project

The Demographic and Health Survey (DHS) Program, funded by the US Agency for International Development (USAID), is contributing unique data to the global database on population and health. To date, over 35 surveys have been conducted in developing countries, including Botswana, Kenya, Nigeria, Senegal, Zimbabwe, Indonesia, Sri Lanka, Peru, Bolivia, Ecuador, Egypt, and Morocco.

The DHS Program, administered by the Institute for Resource Development, Columbia, Md, began in 1984. The primary objectives of the program are to assist developing countries in building the technical skills and resources necessary for conducting demographic and health surveys and to provide developing nations with information for informed health and population policy choices. The data are gathered through a core questionnaire that is adapted to local needs and cultures.

The DHS core questionnaire has two parts: a household schedule for information on all household members and on household amenities and a woman's questionnaire for information on the woman and her children.

The DHS health data are used to estimate levels of neonatal, infant and child mortality; assess treatment practices for major childhood diseases; estimate coverage rates for maternity care and immunization; determine incidence of high-risk births; and assess regional variation in nutritional status.

The data can also be used to examine the relationships between various sociodemographic variables and health variables to better understand the health picture of developing countries.

The DHS survey data, which are maintained in a computerized archive, are available for 28 countries. Usually, several datasets are available for each country. These include women's data, household data, and service availability data.

In general, DHS datasets are available to researchers after the publication of the country survey report. For a few surveys, written permission is necessary from the in-country survey organization. The data files are accompanied by questionnaires, machine-readable descriptions and associated documentation. Currently, the cost of a DHS dataset is $200 ($50 for researchers in developing countries).

To order a DHS dataset, send a completed request form and description of your analysis project to the DHS Data Archive. Request forms are available from the archive. The address is DHS Data Archive, Institute for Resource Development, 8850 Sanford Boulevard, Columbia, MD 21045, USA. Tel: 301/290-2977; Telex: 87775; Fax: 301/290-2999.

Smoking, Social Support, and Hassles in an Urban African-American Community

Patrick S. Romano, MD, MPH, Joan Bloom, PhD, and S. Leonard Syme, PhD

ABSTRACT

Background. Despite public health efforts, the prevalence of smoking among African Americans remains high. The determinants of smoking behavior in this population must be elucidated so that interventions can be better targeted and more effective.

Methods. As part of a prospective community intervention trial to reduce cancer mortality, we conducted a random household survey of 1137 African-American adults in San Francisco and Oakland between November 1985 and July 1986. The survey instrument included questions about social network characteristics, instrumental and emotional aspects of social support, smoking behavior, and stressors.

Results. The overall prevalence of smoking (41.9%) was higher than that reported in national surveys. Logistic models revealed that persons reporting high levels of stress, represented by an abbreviated hassles index, were more likely to smoke than those reporting less stress. Women with poor social networks were more likely to smoke (odds ratio = 3.1) than women with optimal networks; however, this relationship did not hold among men. Indeed, men lacking emotional support from friends or family were less likely to smoke (odds ratio = 0.5) than men receiving such support. No interaction between social support and hassles was observed.

Conclusions. Stressful environments may contribute to high-risk smoking behavior among urban African Americans. (*Am J Public Health.* 1991;81:1415–1422)

Introduction

Educational efforts and innovative policies have reduced the prevalence of cigarette smoking among US adults from 41.7% in 1965[1] to 28.8% in 1987.[2] Despite this progress, 15.7% of deaths in 1984 could still be attributed to smoking.[3] African Americans suffer from a particularly severe burden of smoking-related disease. The age-adjusted smoking-attributable mortality rate among African Americans in 1984 was 143.2 per 100 000 persons compared with 119.0 among Whites.[3] The 1987 National Health Interview Survey found that 32.9% of African-American adults smoked compared with 28.5% of White adults; the proportion of ever smokers who had quit was far lower among African Americans (31.0% vs 45.8%).[2]

Much attention has therefore been directed to reducing the prevalence of smoking among African Americans.[4,5] Identification of factors that promote smoking in the African-American community is a prerequisite to developing effective interventions. Evidence from the 1987 General Social Survey suggests that stress may be such a factor.[6]

Stress and Smoking

Psychological stress can be defined as an "internal subjective state involving the perception of threat to one's well-being," while stressors are stimuli that provoke psychological stress in susceptible individuals.[7] Stressors have been shown to increase tobacco consumption among established smokers,[8] promote adolescent cigarette use,[9] predispose to smoking clinic failure,[10] and precipitate relapses among successful quitters.[11,12] Investigators have evaluated a variety of potential stressors, including anxiety-inducing tasks,[8] family turmoil,[9] job factors,[13] and major life events (e.g., unemployment, divorce).[6,10]

Lazarus and colleagues have recently argued that daily hassles may represent a better conceptualization of stress than traditional measures of major life events.[14] They define hassles as "the irritating, frustrating, distressing demands that to some degree characterize everyday transactions with the environment."[15] Hassles appear to predict many health outcomes, including psychological distress,[15] somatic symptoms,[16] respiratory illness episodes,[17] and arthritis-related disability.[18]

In a 12-month follow-up study of ex-smokers who had participated in a smoking cessation program, the frequency and severity of hassles were significant predictors of late relapse in men, but not in women.[19] A similar evaluation of a work-site smoking cessation program found that the mean severity of hassles, but not the count, predicted relapse by 12 months.[20] Previous studies have not addressed the relationship between hassles and the prevalence of smoking in a community sample. Given the role of smoking as a psychological coping behavior,[7] we hypothesized that persons reporting numerous hassles

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Patrick S. Romano is with the Institute for Health Policy Studies of the University of California, San Francisco. Joan Bloom is with the Department of Social and Administrative Health Sciences, and S. Leonard Syme is with the Department of Behavioral and Environmental Health Sciences of the University of California, Berkeley, School of Public Health.

Requests for reprints should be sent to Patrick S. Romano, MD, UCSF Institute for Health Policy Studies, 1388 Sutter Street, 11th Floor, San Francisco, CA 94109.

This paper was submitted to the journal August 27, 1990, and accepted with revisions April 4, 1991.

TABLE 1—Items from the Abbreviated Hassles Index (n = 1134)

Description	% of Respondents Reporting
Being out of work for a month or longer	47.9
Having a serious illness or accident	28.9[a]
Not having enough money for food, clothing, housing, or other necessities of life	31.0
Being concerned about getting credit	26.2
Having a problem with getting things repaired around the house	29.3
Having a check late or lost in the mail	21.1
Having something stolen or having the house or car broken into	24.9
Having a violent argument with a friend or relative	23.5
Having some other kind of trouble with family members	25.7
Being concerned about living in an unsafe area	36.8

[a] This item was deleted from the hassles index used in multivariate regressions to predict smoking behavior because it might represent a consequence rather than a cause of smoking.

would be more likely to smoke than those experiencing fewer stressors.

Social Support and Smoking

There is extensive epidemiologic evidence of a relationship between social support and diverse health outcomes.[21] Persons with high levels of social support have lower mortality rates,[22] have less coronary atherosclerosis,[23] and use less health care[24] than those with lower levels of support. Social support may exert these salutary effects partially by fostering changes in health-related behaviors, such as smoking. Indeed, successful quitters have higher levels of perceived support from their partners than smokers who never quit or relapse.[25] Having someone to talk to about one's problems predicted cessation in another recent study.[26] The prevalence of smoking is higher among separated or divorced persons than among those currently married or widowed,[2] even after controlling for other demographic characteristics.[27] We therefore hypothesized that persons with poor social support would be more likely to smoke than those with good support.

Buffering and the Stress–Support Interaction

The mechanism by which social support improves health outcomes has been a subject of intense debate.[28] Advocates of the buffering hypothesis argue that social support works by moderating the deleterious effect of stressful events and serving as "coping assistance."[29] Therefore, our final hypothesis was that any association between social support and smoking would be stronger among persons experiencing stressors than among those in low-stressor environments.

The present study was undertaken to assess the cross-sectional relationship between these social factors and smoking behavior in a large population-based sample of African-American adults in the San Francisco Bay area.

Methods

As part of a prospective controlled community intervention trial to reduce cancer mortality, a survey of African-American households in San Francisco and Oakland, Calif, was conducted between November 1985 and July 1986. Potential respondents were identified through two-stage probability sampling. In the first stage, 100 census blocks with at least 25% African-American population (according to the 1980 census) were randomly selected from each city. Specific households were then randomly selected from each block. Using the Kish procedure, interviewers randomly identified one respondent from each household with an African-American member aged 20 years or over.[30] A non-profit survey research organization hired, trained, and supervised field workers. Intensive efforts were made to interview persons who were not available during the initial visit. Follow-up telephone contacts with at least 10% of all respondents confirmed that the interviews took place and that the information obtained was accurate.

The survey instrument included questions about current health status, health behaviors, use of preventive services, sources of health information, knowledge and attitudes regarding cancer, social networks and stressors, and demographic characteristics. Smokers were defined in the usual manner[2] as individuals who currently smoked and had smoked at least 100 cigarettes since birth. Average daily consumption of 20 or more cigarettes was used as the cutoff to separate heavy smokers from light smokers. Demographic variables included age, gender, years of formal education, household income (recoded from an ordinal into a continuous variable), marital status, and employment. Respondents were asked to describe the impact of smoking on cancer risk (1 through 7 scale); those who indicated that it had no effect, a favorable effect, or an unknown effect were categorized as "unaware."

A four-item, health-specific locus of control scale was adapted from the work of Wallston et al.[31] This scale represents the mean of standardized Likert-type responses (coded 1 through 6) to four statements concerning self-control over health. A high score represents a strong belief in one's ability to avoid getting sick by taking care of oneself, to stay healthy by following "the advice of experts," and to control "the most important things that affect health."

A 10-item hassles index was created by abbreviating the original Kanner et al. list of 117 hassles.[15] Each item was chosen to represent a dimension that community residents involved in the project perceived to be especially relevant (Table 1). The second item ("having a serious illness or accident") was deleted from the scale because of concern that such events might be consequences rather than causes of smoking.[32] Scoring from zero to nine was based upon how many of the remaining hassles happened to the respondent or a person "most important" to him or her during the preceding three months. We did not employ an intensity scale because the frequency of hassles has superior predictive power.[15–17] In accord with Dohrenwend's recommendations,[32] we inquired whether these events occurred but did not ask for any judgment as to their significance.

The original Kanner et al. instrument was reported[17] to have a test–retest reliability of $r = 0.79$ with nine consecutive monthly administrations. Similar instruments developed by Weinberger et al.,[18] Wolf et al.,[33] and DeLongis et al.[34] have demonstrated 5- through 9-month median test–retest reliabilities of .72 to .77. Our abbreviated instrument had good internal consistency (Cronbach's $\alpha = .74$). We validated the scale against demographic variables; the mean score was 2.5 for employed and 4.1 for unemployed respondents. Hassle scores were inversely related to family income ($r = -.19$, $P = .0001$).

Measures of social network characteristics can be categorized as either structural or functional. Structural features were assessed using Berkman and Syme's Social Network Index (SNI),[22] which measures four types of social ties: marital status, number of close friends, number of close relatives, and membership in formal groups. This index was modified slightly by adding two types of specified organizations (neighborhood and senior groups) and dropping a question on the frequency of contact with close friends or relatives. Scores range from one to four, where one represents those with the fewest relatives or friends and no history of church or group involvement. Functional features were assessed using Seeman and Syme's three-item Network Instrumental Support scale,[23] which measures whether respondents would "most likely" turn to friends or family for assistance with rides, borrowing "a small sum of money," or household repairs. Their four-item Network Emotional Support scale[23] was abbreviated to two items, only one of which had construct validity ("Who would you turn to for information or advice about a personal problem?"). This question, scored one for those who cited friends or family and zero for others, resembles the emotional support item used by Strogatz and James.[35]

The distribution of demographic characteristics was examined among smoking status categories: never smokers, ex-smokers, current light smokers, and current heavy smokers. Analysis of variance with two-tailed pairwise t tests and Bonferroni's correction for multiple comparisons was used to contrast the means of continuous variables. Chi-square analysis, with Yates' continuity correction for 2 × 2 tables, was performed on categorical variables. The Kruskal-Wallis procedure with two-tailed pairwise mean rank comparisons and Bonferroni's correction was applied to ordinal variables.

Multivariate logistic regression models were constructed to determine whether social support and hassles contributed to explaining smoking status after controlling for potential confounders. Heavy smokers and light smokers were aggregated because of their similar characteristics. In the absence of data about when or why respondents quit, both never smokers and ex-smokers were categorized as non-smokers in our primary analyses. Independent variables included age (with a quadratic term), gender, educational status (<12 years, 12 years, 13 through 15 years, and ≥16 years), household income (transformed to optimize fit), health-specific locus of control, and lack of awareness about the risk of smoking. These variables were selected on the basis of previous studies of smoking correlates.[2,36]

The abbreviated hassles index was then added to the basic model to determine the role of psychosocial stress, controlling for demographic and attitudinal characteristics. Social support variables were added next, with the reference group consisting of individuals with SNI = 4 (optimal) and Network Instrumental Support and emotional support scales greater than zero. To test the hypothesis that social support buffers the deleterious effects of stress, interaction terms between the hassles index and the three support measures were tested. All two-factor interactions involving demographics and social support were evaluated by forward selection and backward elimination procedures.[37] The significance level for entry and exit was $\alpha = .10$, with adjustment to ensure that terms derived from a single variable entered together. Hierarchical models were compared with the likelihood ratio test.

Finally, multivariate linear regression was used to determine factors associated with self-reported daily cigarette consumption among current smokers. Multicollinearities prohibited inclusion of interaction terms in this model. The PROC LOGIST and PROC REG programs in SAS were used for most analyses.[38] Odds ratios (ORs) and 95% confidence intervals (CIs) were calculated by exponentiating the logistic coefficient estimates. These ORs modestly overestimate the true risk ratios because the "rare disease" assumption underlying the OR is not satisfied.

Results

The total sample size was 1137, representing a response rate of 67.6% in Oakland and 69.1% in San Francisco. The age, gender, and income distributions of our sample closely resembled 1980 census data from the same neighborhoods. The San Francisco and Oakland samples were almost equal in size (569 and 568, respectively) and showed similar demographic characteristics. Therefore, they have been combined for all analyses.

Of the 1134 respondents who described their smoking habits, 36.0% never smoked, 22.1% had quit smoking, 26.8% were light smokers, and 15.1% were heavy smokers. Current smoking was more prevalent among women than among men (43.1% vs 40.1%), but female smokers were less likely than male smokers to consume 20 or more cigarettes daily.

Sociodemographic Characteristics by Smoking Status

Table 2 illustrates how the sociodemographic characteristics of current smokers differed from those of ex-smokers and never smokers. Ex-smokers were significantly older than current or never smokers. Ex-smokers were more than twice as likely to have graduated from college as heavy smokers (15.6% vs 7.1%). About 41% of heavy smokers and 36% of light smokers were separated or divorced, compared with only 20% of never smokers and 24% of ex-smokers. Current smokers reported lower household incomes than ex-smokers. Heavy smokers had a particularly high frequency of disability (15.4%); both light and heavy smokers were more likely to be unemployed than ex-smokers or never smokers. Current smokers had higher hassle scores than either ex-smokers or never smokers, whereas the latter two groups had higher SNI scores. More than 15% of heavy smokers denied a link between smoking and cancer, as compared with 4% of nonsmokers.

Intercorrelations Between Stress and Social Support Measures

Bivariate Pearson correlations between the social support and stress measures revealed only two significant associations. The hassles score was negatively related to the SNI ($r = -.10$, $P = .0009$), whereas Network Instrumental Support was positively related to emotional support ($r = .17$, $P = .0001$). The structural and functional measures of social support were not related to each other.

Demographic Predictors of Smoking Status

Table 3 shows the beta coefficients and ORs with 95% confidence limits for each variable in the basic logistic model, which included demographics, risk awareness, and locus of control. Lack of awareness of the smoking–cancer link was a strong correlate of current smoking, whereas health-specific locus of control

TABLE 2—Distribution of Selected Sociodemographic Characteristics by Smoking Status

Characteristics	Never Smokers (n = 406)	Ex-Smokers (n = 251)	Light Smokers (n = 304)	Heavy Smokers (n = 171)
Mean age, y (F = 15.14, $P < .0001$)	44.3[a]	48.7[b]	39.8	42.0
Gender (χ^2 = 31.72, $P < .0001$), %				
Male	34.3	55.4	36.8	43.9
Education (χ^2 = 21.47, $P = .01$), %				
Less than high school	26.9	27.2	20.2	30.2
High school graduate or equivalent	27.6	29.2	38.4	31.4
1–3 y college	33.8	28.0	31.8	31.4
4 or more y college	11.7	15.6	9.6	7.1
Marital status (χ^2 = 77.16, $P < .0001$), %				
Married	36.2	49.0	25.3	28.1
Separated/divorced	20.4	24.5	36.5	41.5
Widowed	16.5	10.8	8.2	9.4
Never married	26.9	15.7	29.9	21.0
Mean annual household income, dollars (F = 5.89, $P < .001$)	20 200	22 502[c]	17 663	18 164
Perceived risk of cancer from smoking, % (χ^2 = 32.16, $P < .001$)				
Not aware	3.7	4.0	10.3	15.4
Employment (χ^2 = 43.59, $P < .0001$), %				
Full or part time	57.6	54.8	57.9	50.3
Unemployed	8.1	10.4	14.5	16.6
Disabled/ill	5.4	4.4	5.6	15.4
Not in labor force	28.8	30.4	22.0	17.7
Locus of control, mean score (F = 1.21, $P = .30$)	0.03	0.04	−0.03	−0.05
Social Network Index, mean score (range 1–4, χ^2 = 33.67, $P < .0001$)	2.47[d]	2.62[e]	2.29	2.07
Network Instrumental Support, mean score (range 0–3, χ^2 = 7.38, $P = .06$)	1.51	1.34	1.51	1.55
Hassles Index, mean score (range 0–9, χ^2 = 22.81, $P < .0001$)	2.37[f]	2.37[f]	2.95	3.28

[a] Different from ex-smokers ($P < .05$) and light smokers ($P < .01$).
[b] Different from never smokers ($P < .05$), light smokers ($P < .001$), and heavy smokers ($P < .01$).
[c] Different from light smokers ($P < .01$) and heavy smokers ($P < .05$).
[d] Different from heavy smokers ($P < .01$).
[e] Different from light smokers ($P < .05$) and heavy smokers ($P < .001$).
[f] Different from light smokers ($P < .05$) and heavy smokers ($P < .05$).

was marginally significant. The quadratic coefficient for age indicates that smoking prevalence peaked in middle age, with lower rates in young and elderly adults. College graduates were less likely to smoke than high school graduates, but no gender difference was noted. Current smoking was inversely related to household income.

Psychosocial Predictors of Smoking Status

When we added the abbreviated hassles index alone to this model, it significantly contributed to explaining smoking risk (χ^2 = 5.54, df = 1, P = .02). We next added the five social support variables (SNI levels, Network Instrumental Support, and emotional support) and four interaction terms (Table 4). Only gender interactions with the SNI and emotional support scales met our entry criteria. Interaction terms between the hassles index and social support variables were not significant, either individually or as a group (χ^2 = 5.82, df = 5, P = .32). The model in Table 4, which includes social network characteristics, was superior to that with only demographics and hassles (χ^2 = 30.68, df = 9, P = .0001) in predicting smoking status.

A high frequency of hassles in the preceding 3 months was associated with increased likelihood of smoking for both sexes. With an OR of 1.07 per unit hassle, an individual with five hassles had 1.43 times the odds of smoking as a similar individual with no hassles. Strong social networks were associated with lower odds of smoking among women, but not among men. Conversely, men with inadequate emotional support were *less* likely to smoke than men with better emotional support. There was no relationship between this item and smoking among women. Instrumental support was independent of smoking.

Confirmatory Models

The robustness of these findings was tested in several ways. By using an income variable in regression equations, we lost 85 observations for which we had otherwise complete data. To examine the effect of this loss, we tested a model with employment status (employed, unemployed, not in labor force) in place of income. The prevalence of smoking was highest among those not in the labor force (OR = 1.48; CI = 1.03, 2.13), and the coefficients for psychosocial variables remained significant. Income and employment could not be included in the same model because of collinearity. We also tested a model in which social support variables were entered in continuous form. The findings were similar to those in Table 4, except that the interaction between gen-

TABLE 3—Estimated Logistic Regression Coefficients and Odds Ratios Comparing Current Smokers with Nonsmokers in Model 1 (demographic variables only)

Variable	Coefficient	SE	Adjusted Odds Ratio	95% Confidence Interval
Intercept	−1.324**	.639		
Locus of control (mean of 4 standardized items)	−0.210*	.114	0.81[a]	0.65, 1.01
Lack of risk awareness (not aware of cancer risk = 1)	1.782***	.315	5.9	3.2, 11.0
Age (y)	0.116***	.031		
Age^2 (y^2)	−0.0016***	.0003		
Gender (male = 1)	0.064	.140	1.07	0.81, 1.40
Education				
Less than high school	−0.094	.199	0.91	0.62, 1.35
High school graduate or equivalent	Referent			
1–3 y college	−0.262	.169	0.77	0.55, 1.07
4 or more y college	−0.624**	.249	0.54	0.33, 0.87
Annual household income (thousands, square root)	−0.163***	.047	0.85[b]	0.78, 0.93

Note. The dependent variable is current smoking. n = 1108; model χ^2 = 114.22 (*df* = 9); −2logL = 1263.09.

[a] Adjusted odds ratio associated with a change of one standard deviation in the locus of control scale.

[b] Adjusted odds ratio associated with a one-unit change in the transformed income variable (e.g., $16 000 per year vs $9000 per year).

**P* < .10.

***P* < .05.

****P* < .01.

der and SNI was weakened by implicitly forcing it to be uniform across SNI levels (χ^2 = 1.86, *df* = 1, *P* = .17). A multiple linear regression model was used to test whether the association between psychosocial stress and smoking was driven by a cohort of light, relatively low-risk smokers. Hassle scores were positively related to self-reported cigarette consumption among current smokers, after adjusting for demographic characteristics (*t* = 1.76, *P* = .08). Social support variables did not improve the model (F = 1.05, *df* = 5/407, *P* = .39).

Since marital status is a component of the SNI, we examined whether the effect of that measure could be explained simply by a higher prevalence of smoking in separated and divorced adults. A logistic regression model that included SNI and marital status was significantly superior (χ^2 = 12.53, *df* = 6, *P* = .05) to an otherwise identical model without SNI. Stratified analyses confirmed that lack of church involvement was related to smoking among married, separated/divorced, and widowed respondents. A small number of friends and nonmembership in community organizations were related to smoking among married persons, but not in other marital categories.

To explore whether psychosocial factors were related primarily to initiation or cessation of smoking, we performed a logistic analysis similar to the one in Table 4 but limited to ever smokers, with quitting as the dependent variable. Low SNI scores were associated with low odds of having quit smoking among women, but the hassles index was independent of smoking cessation (χ^2 = .06, *df* = 1, *P* = .81). Age appeared to be the major confounder of the bivariate association between hassles and quitting (Table 2).

Discussion

The prevalence of smoking in this population-based probability sample of African-American adults was higher than that reported in studies using similar sampling techniques in the general population.[2] For example, the 1985 Current Population Survey found that 26% of Californians 20 years or older currently smoke.[39] The 1985 Behavioral Risk Factor Survey revealed current smoking rates of 26.3% and 24.9% in male and female Californians, respectively.[40] The gap between these rates and that observed by our group (41.9%) suggests that African Americans living in urban communities with large African-American populations are at especially high risk for smoking-related illness. Aggressive marketing by tobacco companies in these neighborhoods may be a contributing factor.[41]

Our study of African-American adults found that the prevalence of current smoking was highest among low-income, middle-aged individuals who never completed college. These findings are similar to those reported from the National Health Interview Survey[2] and a recent survey of Black insurance policyholders.[42] Our finding that individuals who lack awareness of the risks of smoking are more likely to smoke also corroborates previous work, particularly with adolescents.[43] This association may be explained in two ways: (1) People who are ignorant of the risks of smoking misperceive their vulnerability, or (2) nicotine addicts deny those risks to minimize cognitive dissonance.

Our measure of health-specific locus of control was only marginally related to smoking in multivariate models; however, individuals with a strong sense of control over health determinants tended to be less likely to smoke. Since the relationship between locus of control and smoking behavior has not been well studied, this finding should stimulate further investigation. The four-item scale we employed may not adequately represent the concept.

The independent association between hassles and current smoking has not previously been recognized. This effect was uniform by gender and did not change significantly after including social networks as a potential confounder. Among current smokers, hassled individuals tended to smoke more cigarettes daily than those relatively free of hassles. These findings support Lazarus' theory that the stresses of everyday life exert a cumulative impact on health.[14] Smoking may rep-

TABLE 4—Estimated Logistic Regression Coefficients and Odds Ratios Comparing Current Smokers with Nonsmokers in Model 2 (demographic, stress, and social support variables with interactions)

Variable	Coefficient	SE	Adjusted Odds Ratio	95% Confidence Interval
Intercept	−2.770***	.726		
Locus of control (mean of 4 standardized items)	−0.200*	.116	.82[a]	0.65, 1.03
Lack of risk awareness (not aware of cancer risk = 1)	1.752***	.320	5.8	3.1, 10.8
Age (y)	0.125***	.031		
Age^2 (y^2)	−0.0016***	.0003		
Gender (male = 1)	0.847**	.341	[b]	[b]
Education				
Less than high school	−0.112	.206	0.89	0.60, 1.34
High school graduate or equivalent	Referent			
1–3 y college	−0.244	.173	0.78	0.56, 1.10
4 or more y college	−0.585**	.258	0.56	0.34, 0.92
Annual household income (thousands, square root)	−0.111**	.050	0.89[c]	0.82, 0.99
Social Network Index (SNI)				
Level 1 (smallest)	1.123***	.314	3.1[d]	1.7, 5.7
Level 2	0.715***	.256	2.0[d]	1.2, 3.3
Level 3	0.871**	.349	2.4[d]	1.2, 4.7
Level 4 (largest)	Referent			
Network Instrumental Support (none from family or friends = 1)	−0.186	.196	0.83	0.56, 1.22
Network Emotional Support (none from family or friends = 1)	0.059	.185	1.06[d]	0.74, 1.52
Hassles Index (1–9 scale)	0.072**	.032	1.07[e]	1.01, 1.14
Gender Interactions (male = 1)				
SNI Level 1	−0.920**	.442	1.22[f]	0.66, 2.29
SNI Level 2	−0.112	.373	1.83[f]	1.07, 3.13
SNI Level 3	−0.794*	.432	1.08[f]	0.56, 2.08
Network Emotional Support	−0.742***	.281	0.51[f]	0.33, 0.77

Note. The dependent variable is current smoking. $n = 1008$; model $\chi^2 = 138.89$ ($df = 18$); $-2\log L = 1226.87$.

[a] Adjusted odds ratio associated with a change of one standard deviation in the locus of control scale.

[b] Adjusted odds ratios for male gender vary according to the level of emotional support and social network involvement, from 0.44 (0.22, 0.88) among those with poor emotional support and SNI = 1 to 2.33 (1.19, 4.56) among those with optimal levels of both (SNI = 4). These odds ratios were derived by exponentiating the sum of the gender and applicable social support interaction coefficients.

[c] Adjusted odds ratio associated with a one-unit change in the transformed income variable (e.g., \$16 000 per year versus \$8000 per year).

[d] Adjusted odds ratios for women with low SNI scores (1, 2, or 3) or absent emotional support, relative to women with SNI = 4 and appropriate emotional support, respectively.

[e] Adjusted odds ratio associated with a one-unit increase in the hassles score.

[f] Adjusted odds ratios for men with low SNI scores (1, 2, or 3) or absent emotional support, relative to men with SNI = 4 and appropriate emotional support, respectively. These odds ratios were derived by exponentiating the sum of the social support and applicable gender interaction coefficients.

*$P < .10$.
**$P < .05$.
***$P < .01$.

resent a coping behavior that mitigates the harmful psychological impact of a stressful environment. After controlling for other variables, hassles were not associated with ex-smoking among ever smokers. This finding suggests that hassles may influence smoking initiation more than cessation, but the cross-sectional nature of our data undercuts such an interpretation.

We observed that social support, as measured by the SNI, was inversely related to the likelihood of smoking only among women. This gender interaction may be a consequence of our measure or an inherent feature of how social support affects health behaviors. Male respondents may have exaggerated their social networks to appear well-adjusted to female interviewers. Men may be more responsive to unmeasured dimensions of social support, such as peer support directly related to quitting. Waldron and Lye[27] demonstrated that marital status is related to current smoking among both sexes, but the other three SNI components may be irrelevant for men. The potentially favorable effects of social networks for men may be outweighed by concomitant peer pressure to smoke. This hypothesis is supported by the observation that men who turn to close friends or relatives for help with personal problems are more likely to smoke than self-reliant men.

These data do not support the hypothesis that social support buffers the adverse impact of stressful events.[28] Instead, we found that social networks were associated with decreased smoking in women at all levels of hassles. Conversely, strong social support did not attenuate the detrimental effect of hassles. Social networks did not appear to operate by enhancing instrumental or emotional support for our female respondents.

Household interview surveys are susceptible to several sources of bias. Our sampling scheme excluded African-American families residing in predominantly non-Black census blocks. The further exclusion of institutionalized individuals and the homeless may limit the generalizability of our findings to the entire African-American community. Misclassification bias is always a concern in measuring sensitive behaviors, such as smoking. However, nondifferential misclassification of smokers as nonsmokers would not create artifactual associations between smoking and social factors.

Since our study design was cross-sectional, we cannot elucidate the chain of causation that links social support, stress, and smoking. Given present public attitudes toward tobacco, smokers who quit may experience improved social connections. Similarly, smokers may be more likely to suffer hassles than nonsmokers as a result of covert discrimination or physiological withdrawal symptoms. Recall bias may explain the associations if smokers reported more hassles or less social support in a post hoc effort to justify their unhealthy behavior. The likelihood of biased responses was minimized by separating questions about smoking from those about social factors and by disguising the research hypotheses. Finally, hassled persons and women with low SNI scores may possess other characteristics that are the actual determinants of smoking. The SNI could be a proxy for future orientation or interest in self-improvement; perceived hassles could reflect an underlying psychological vulnerability that predisposes to smoking.

The high prevalence of smoking noted among urban African Americans underscores the importance of smoking prevention and cessation activities focused on this population. Prospective studies with more detailed measures of support and stress are needed to clarify the direction of causation between psychosocial factors and smoking. The generalizability of our findings to other ethnic groups should also be evaluated. To the extent that smoking represents a behavioral response to stressful circumstances, prevention and cessation programs may achieve optimal success only by addressing larger community problems. As an interim measure, interventions to teach healthier coping behaviors may be efficacious. Interventions designed to improve social networks may have a particular role in smoking cessation programs for African-American women. ☐

Acknowledgments

This research was supported by Grant CA 40651 from the National Cancer Institute. General support for the first author was provided through an Institutional National Research Service Award (HS00026) from the Agency for Health Care Policy and Research. The efforts of our project director, William A. Hayes, PhD, are gratefully acknowledged.

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Preventive Health Care for Children Works Better in Europe

Although they spend a smaller share of their gross national product on health care than does the United States, European countries achieve nearly complete participation of children in preventive health care and more favorable health outcomes for young children and adolescents. These are among the findings of a just-published report, *Preventive Health Care for Young Children: Findings from a 10-Country Study and Directions for United States Policy,* by Bret C. Williams, MD, MPH and C. Arden Miller, MD.

Supported by the William T. Grant Foundation and the US Maternal and Child Health Bureau, Williams and Miller looked at preventive health services for children from infancy through adolescence and at the social benefit programs that support children and families in Belgium, Denmark, France, the Federal Republic of Germany, Ireland, the Netherlands, Norway, Spain, Switzerland, and the United Kingdom. They found the following:

- Childhood mortality rates are higher in the United States than in Europe at all age levels, most markedly among children aged 1 to 4 years and adolescents aged 15 to 19 years.
- US immunization rates among preschool children lag behind European rates by 23% to 49%.
- US children are less likely to have a regular source of medical care than their European counterparts.
- US children are far more likely than children in Europe to experience injury-related deaths.

Williams and Miller found that the "truly remarkable aspect" of preventive health services in Europe is the nearly complete participation of children. Compliance with the full schedule of preventive health care visits generally exceeds 90% during the first year of life; completed immunization rates among young children are also generally above 90%. All health care for children in the study countries is financed under national systems that assure financial coverage of the entire population without means testing. Williams and Miller note that full availability of health care for children and social benefits programs for families came about in Europe only after earlier selective approaches failed.

Preventive Health Care for Young Children suggests that six attributes of European health systems are particularly worthy of US consideration:

1. Access to one or more provider systems that assure compliance with routine preventive services and linkage to more sophisticated care when necessary;
2. Separation of preventive services from the delivery of acute medical care, allowing the former to be rendered by less specialized personnel;
3. No means testing or payment required at the time of service delivery;
4. Linkage of families with children to benefit programs that alleviate poverty equitably across age groups;
5. Access of very young children to low-cost regulated programs for children of working parents;
6. A tracking system for children from birth through the age of 4 years to assure that young children and their families actively participate in health and social programs for which they are eligible.

Williams and Miller point out that European patterns of preventive health care and social supports are not strikingly different from the proposals of professional groups and public health agencies in the United States, including the majority of members of the National Commission on Children. They suggest that a national effort to improve preventive health care for children in the United States could build on what is being learned from US initiatives that include local prenatal health care programs, statewide tracking systems, and established, but underfunded, national efforts like WIC and Head Start.

Preventive Health Care for Young Children is a publication of the National Center for Clinical Infant Programs (NCCIP), a national nonprofit organization concerned with the healthy development of children and families during the first three years of life. In 1987, NCCIP published C. Arden Miller's monograph, *Maternal Health and Infant Survival,* an analysis of medical and social services to pregnant women, newborns, and their families in the same 10 European countries described in the new report.

Copies of *Preventive Health Care for Young Children* are available for $9.50 each; *Maternal Health and Infant Survival* is available for $7.50 per copy. Both reports may be ordered as a set for $15.00. Checks should be made payable to NCCIP and include shipping and handling charges of $3 for the first item ordered, $1 for each additional item, and 10% for orders of $60 or more. Send orders to the National Center for Clinical Infant Programs, PO Box 25494, Richmond, VA 23260-3494.

Cardiovascular Risk Factors among Asian Americans Living in Northern California

ABSTRACT

Background. Recent substantial immigration has created large population subsets of Asian Americans in the United States. Available data about cardiovascular risk factors in these persons are sparse.

Methods. This study examined data among 13 031 persons self-classified as 5951 Chinese, 4211 Filipinos, 1703 Japanese, and 1166 other Asians. Covariates in regression analyses were age, smoking, alcohol, education, and marital status.

Results. Chinese men and women had the lowest adjusted mean body mass index. Filipino men and women had the highest prevalence of hypertension. There were no major differences in blood glucose levels. Total cholesterol levels were highest in Japanese men and women. Comparisons of US-born persons and those born in respective countries of origin showed no major cholesterol or glucose differences; more hypertension only in Chinese and other Asian men; higher body mass index in men, but not in women of most ethnicities; and a lower smoking prevalence in men, but a substantially higher one in women.

Conclusions. These data show important ethnic differences in risk factors among Asian Americans and indicate groups that should be targeted for public health efforts concerned with obesity (Asian-American men), hypertension (Filipino-American men and women), hypercholesteremia (all Asian Americans), and smoking cessation (Asian-American women). (*Am J Public Health.* 1991;81:1423–1428)

Arthur L. Klatsky, MD, and Mary Anne Armstrong, MA

Introduction

Despite a marked acceleration of US immigration from Asia since the 1960s,[1,2] there are few reports concerning cardiovascular risk factors and disease incidence in Asian Americans. The available data[3–7] suggest more coronary artery disease among US-resident Japanese Americans (vs those resident in Japan), related to life-style changes. A study of elderly Chinese Americans[8] suggests risk factor profiles similar to those in China, with low risk of coronary artery disease and high risk of cerebrovascular episodes.[9,10] Two reports concerning blood pressure among Asian Americans in California[11,12] suggest that Filipinos are at greater risk; the authors of one[11] suggest that sociocultural factors are involved, but the authors of the other[12] raise the possibility of genetic influences. We have been given an opportunity to study cardiovascular risk factors among a substantial number of free-living Asian Americans and here present data concerning several of these traits.

Methods

Data were supplied by 13 031 Asian Americans who voluntarily took health examinations offered by a northern California prepaid health care program from the years 1978 through 1985. Sociodemographic and medical history data were obtained from a self-administered questionnaire that included the queries "What is your race?" and "Where were you born?" Previous descriptions have been published.[12,13] Sex and ethnic distributions are included in Table 1.

Examinees were asked to take no food or nutritious drink for at least 4 hours before their appointment time. Supine blood pressure was determined by a Roche Arteriosonde Blood Pressure Monitor, Model 1216. Blood chemistry tests were performed with a Technicon SMA 12-6 Multichannel Analyzer from January through May 1978; thereafter, an American Monitor ICDA Automated Chemistry Analyzer was used.

Adjusted means for body mass index, systolic blood pressure, diastolic blood pressure, total cholesterol, and glucose were determined by multiple regression analysis with the following covariables: age, body mass index (except when the dependent variable was body mass index), marital status, education, and alcohol use. Using the largest subgroup (Chinese) as the reference and with the same covariables in logistic regression models, odds ratios (ORs) were determined for high body mass index ($\geq$ 24.4 kg/m^2 or 3.5 lb/in^2), hypertension ($\geq$ 140/90 mm Hg or on treatment), high cholesterol ($\geq$ 6.21 mmol/L or 240 mg/dL), high glucose ($\geq$ 6.11 mmol/L or 110 mgm/dL), current cigarette smoking, and current cigarette smoking of one pack per day or more. Similarly, but with birthplace as an additional covariable, the relations of various traits to each of these

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Arthur L. Klatsky is with the Division of Cardiology of the Department of Medicine, Kaiser Permanente Medical Center, and Mary Anne Armstrong is with the Division of Research, Kaiser Permanente Medical Care Program, Oakland, Calif.

Requests for reprints should be sent to Arthur L. Klatsky, MD, Division of Cardiology, Kaiser Permanente Medical Center, 280 W MacArthur Boulevard, Oakland, CA 94611.

This paper was submitted to the journal November 27, 1990, and accepted with revisions April 2, 1991.

Editor's Note. See related editorial by Yu on page 1391 of this issue.

TABLE 1—Traits of Persons in Asian-American Ethnic Groups by Sex

Trait	Chinese	Filipino	Japanese	Other Asian
No. (% of total)				
M	2754 (48.2)	1729 (30.3)	688 (12.0)	543 (9.5)
F	3197 (43.7)	2482 (33.9)	1015 (13.9)	623 (8.5)
Age (mean ± SD)				
M	40.2 ± 14.2	39.2 ± 13.4	42.1 ± 14.1	35.9 ± 10.4
F	38.0 ± 13.6	36.8 ± 11.6	40.9 ± 13.4	34.1 ± 10.5
Born in continental United States (%)				
M	39.9	8.3	74.6	8.4
F	41.8	8.7	67.4	12.1
Current smoker (%)				
M	16.2	32.9	22.7	30.9
F	7.3	11.4	18.6	12.6
Smoke ≥ 1 pack/day (%)				
M	4.1	7.1	8.2	6.7
F	1.3	1.7	4.6	1.6
Drank any alcohol in past year (%)				
M	72.9	81.7	87.4	80.6
F	56.1	47.4	77.0	53.0
College graduate (%)				
M	53.1	49.6	59.3	58.4
F	43.7	53.7	46.8	38.8
Married (%)				
M	68.7	78.9	68.4	73.4
F	62.3	68.9	61.9	73.7
Cholesterol ≥ 6.21 mmol/L (%)				
M	26.6	29.8	36.3	20.4
F	20.2	20.6	30.3	13.5
Glucose ≥ 6.11 mmol/L (%)				
M	15.9	16.1	20.8	11.6
F	10.2	10.8	13.2	9.6
Body mass index ≥ 24.4 kg/m^2				
M	26.9	41.8	38.0	28.9
F	12.8	25.5	18.0	14.6
Blood pressure ≥ 140/90 mm Hg[a] (%)				
M	10.5	12.2	12.9	6.2
F	7.8	9.6	8.8	5.4

[a] Or receiving treatment for hypertension.

end points were studied within each Asian-American ethnic subset using country of ethnic origin as the reference group. Thus, for Chinese Americans, those born in Hong Kong (n = 607), Taiwan (n = 245), Hawaii (n = 121), the continental United States (n = 2431), and other places (n = 438) were compared with those born in mainland China (n = 2109). For Japanese Americans, those born in Hawaii (n = 167) and the continental United States (n = 1174) were compared with those born in Japan (n = 331). For Filipino Americans, those born in the United States (n = 352) were compared with those born in the Philippines (n = 3815). The largest subsets of other Asian Americans were 295 persons born in Korea, 178 persons born in Vietnam, 252 persons born in India, 55 persons born in Thailand, and 58 persons born in Pakistan; we compared all 102 born in the United States with those born in all "other Asian" countries combined.

Results

Adiposity

Adjusted mean body mass index was lowest in Chinese men and women and highest in Filipino men and women (Tables 2 and 3); these means translated into higher prevalence (Table 1) and adjusted ORs (Table 4) of obesity for the Japanese, Filipino, and other Asian subsets vs the Chinese reference group. Birthplace in the continental United States was associated with increased ORs of obesity in men of all ethnic subsets (Table 5); this was also true for Chinese men born in Hawaii (OR vs birthplace in China = 2.4; 95% confidence interval [CI] = 1.3, 4.5) and for Japanese men born in Hawaii (OR vs birthplace in Japan = 3.2; 95% CI = 1.5, 6.6). Among women, only other Asians of US birthplace had similar increases in ORs of obesity (Table 5). Hawaiian-born Japanese women had an increased OR of obesity (OR vs those born in Japan = 2.4; 95% CI = 1.3, 4.4), but this was less true for Chinese women born in Hawaii (OR vs birthplace in China = 1.5, 95% CI = 0.8, 2.9). Chinese men and women born in Hong Kong or Taiwan had no substantially different ORs of obesity when compared with their counterparts born in China (data not shown).

Among men and women, age and currently married status were positively related to obesity. Educational status and alcohol consumption were not related to this trait among men but showed inverse relations to obesity in women.

Hypertension

The increased hypertension risk of Filipino Americans has been described.[12] Only in Chinese men was continental US birthplace associated with more hypertension (Table 5). Chinese and Japanese men and women born in Hawaii did not have increased risk of hypertension compared with their counterparts born in China, nor did Chinese men born in Hong Kong or Taiwan, or Chinese women born in Hong Kong. Chinese women born in Hong Kong had *less* risk of hypertension (OR vs China-born = 0.3; 95% CI = 0.1, 0.8).

Cholesterol

In each sex, the adjusted means were lowest in other Asians and higher in Japanese men and women than in other ethnic groups (except for small differences between Japanese and Filipino men) (Tables 2 and 3). Compared with the Chinese reference group, only Japanese men and women had increased ORs for high cholesterol (Table 4). Birthplace in the continental United States (Table 5) or in Hawaii (data not shown) vs birthplace in country of origin carried no substantially different ORs for high cholesterol. The only birthplace subset with such a difference was the Chinese women born in Hong Kong (OR vs birthplace in China = 1.5, 95% CI = 1.1, 2.3).

For comparison, unadjusted mean total serum cholesterol was 5.49 mmol/L (212.4 mg/dL) among Black persons (n = 32 890), 5.57 mmol/L (215.4 mg/dL) among White persons (n = 73 920), and

TABLE 2—Adjusted Mean Risk Factors among Asian-American Men

	Ethnicity (Adjusted mean ± SE)				
Risk Factor	Chinese	Filipino	Japanese	Other Asian	Pairs for Which *P* < .01
Body mass index (kg/m^2) (Quetelet's)	22.9 ± 0.1	23.9 ± 0.1	23.7 ± 0.1	23.6 ± 0.1	C vs all
Systolic blood pressure (mm Hg)	122.2 ± 0.3	124.4 ± 0.4	121.3 ± 0.6	120.3 ± 0.7	F vs all
Diastolic blood pressure (mm Hg)	75.5 ± 0.2	75.5 ± 0.2	73.9 ± 0.4	74.0 ± 0.4	F vs J, O
Total cholesterol (mmol/L)	5.59 ± 0.02	5.67 ± 0.03	5.72 ± 0.04	5.52 ± 0.05	F vs O
Glucose (mmol/L)	5.45 ± 0.03	5.54 ± 0.03	5.54 ± 0.05	5.54 ± 0.06	0

Note. Each risk factor was considered separately as a dependent variable in a linear regression analysis that produced means for each ethnic group, adjusted for age, marital status, education, alcohol intake, and body mass index (except when body mass index was the dependent variable). C = Chinese; F = Filipino; J = Japanese; O = other Asian.

TABLE 3—Adjusted Mean Risk Factors among Asian-American Women

	Ethnicity (adjusted mean ± SE)				
Risk Factor	Chinese	Filipino	Japanese	Other Asian	Pairs for Which *P* < .01
Body mass index (kg/m^2) (Quetelet's)	21.2 ± 0.1	22.6 ± 0.1	21.6 ± 0.1	22.2 ± 0.1	C vs all
Systolic blood pressure (mm Hg)	116.3 ± 0.3	119.2 ± 0.3	115.7 ± 0.5	115.9 ± 0.6	F vs all
Diastolic blood pressure (mm Hg)	71.7 ± 0.2	72.9 ± 0.2	70.3 ± 0.3	70.8 ± 0.4	F vs all
Total cholesterol (mmol/L)	5.33 ± 0.02	5.35 ± 0.02	5.52 ± 0.03	5.26 ± 0.04	J vs all
Glucose (mmol/L)	5.22 ± 0.02	5.24 ± 0.02	5.26 ± 0.04	5.28 ± 0.05	0

Note. Each risk factor was considered separately as a dependent variable in a linear regression analysis that produced means for each ethnic group, adjusted for age, marital status, education, alcohol intake, and body mass index (except when body mass index was the dependent variable). C = Chinese; F = Filipino; J = Japanese; O = other Asian.

5.47 mmol/L (211.6 mg/dL) among all Asian-American persons combined.

Glucose

Study of adjusted mean blood glucose levels showed no major differences in any of the comparisons.

Smoking

Chinese men and women were least likely to smoke any amount or to smoke heavily (Tables 1 and 4). Among women, the Japanese were most likely to be smokers; among men (other than Chinese), the differences in smoking were smaller. Generally, men born in the United States were less likely to smoke than their counterparts born in Asia, although there was little difference with respect to heavier smoking (Table 5). Chinese men born in Hawaii, Hong Kong, or Taiwan were slightly less likely to smoke at all or to smoke heavily compared with Chinese men born in China (data not shown); Japanese men born in Hawaii were slightly less likely to smoke at all and substantially less likely to smoke one pack or more per day (OR = 0.3, 95% CI = 0.1, 0.9) than their counterparts born in Japan.

With the exception of Japanese women, women born in the United States were more likely to smoke and had substantially increased adjusted ORs for heavier smoking (Table 5). Chinese women born in Hawaii (but not those born in Hong Kong or Taiwan) were also more likely to smoke (OR vs Chinese birthplace = 3.6, 95% CI = 1.7, 5.5), or to smoke one pack or more per day (OR = 9.8, 95% CI = 2.1, 46.0). Hawaiian-born Japanese women, like those born in the continental United States, were, compared with their Japan-born counterparts, less likely to smoke any amount (OR = 0.5, 95% CI = 0.2, 0.98) or to smoke one pack or more per day (OR = 0.5, 95% CI = 0.1, 2.3).

Among both sexes, alcohol drinking was strongly related and college graduation inversely related to smoking. Other traits related to smoking among men were age (weak negative), body mass index (weak negative), and married status (strong positive); among women, age (weak negative) and marital status (weak negative) were related to smoking.

Discussion

There is a consensus that coronary artery disease is related substantially to all of the risk factors discussed in this article and that cerebrovascular disease is related to several. Cigarette smoking is presumably not primarily determined genetically, whereas the other risk factors studied are determined by both genetic and life-style factors. The presumption that US-born persons of each Asian-American ethnic

TABLE 4—Odds Ratios for Risk Factors among Asian-American Ethnic Groups

Risk Factor	Ethnicity: Chinese[a] OR	Filipino OR (95% CI)	Japanese OR (95% CI)	Other Asian OR (95% CI)
Body mass index ≥ 24.4 kg/m²				
M	1.0	1.9 (1.7, 2.3)	1.6 (1.3, 1.9)	1.8 (1.5, 2.2)
F	1.0	2.7 (2.3, 3.2)	1.4 (1.2, 1.8)	2.0 (1.6, 2.6)
Hypertension[b]				
M	1.0	1.3 (1.0, 1.6)	1.1 (0.8, 1.4)	0.9 (0.6, 1.3)
F	1.0	1.5 (1.2, 1.9)	0.9 (0.7, 1.3)	1.0 (0.6, 1.5)
Cholesterol ≥ 6.21 mmol/L				
M	1.0	1.1 (1.0, 1.3)	1.3 (1.1, 1.6)	0.9 (0.7, 1.2)
F	1.0	1.1 (1.0, 1.3)	1.5 (1.2, 1.8)	1.0 (0.7, 1.2)
Glucose ≥ 6.11 mmol/L				
M	1.0	1.1 (0.9, 1.3)	1.3 (1.0, 1.6)	1.1 (0.8, 1.4)
F	1.0	1.1 (0.9, 1.3)	1.2 (0.9, 1.5)	1.4 (1.1, 1.9)
Smoke any amount (current)				
M	1.0	2.3 (2.0, 2.7)	1.4 (1.1, 1.8)	2.0 (1.6, 2.6)
F	1.0	1.9 (1.6, 2.3)	2.3 (1.9, 2.9)	2.0 (1.5, 2.7)
Smoke ≥ 1 pack/day				
M	1.0	1.6 (1.2, 2.1)	1.8 (1.3, 2.6)	1.7 (1.1, 2.5)
F	1.0	1.5 (0.9, 2.3)	2.8 (1.8, 4.4)	1.5 (0.8, 2.9)

Note. OR = odds ratio; CI = confidence interval. ORs were computed from coefficients estimated by logistic regression for each risk factor and each gender separately, controlled for age, marital status, education, alcohol intake, and body mass index (except when body mass index ≥ 24.4 kg/m² was the dependent variable).

[a] Chinese are reference category.

[b] ≥ 140/90 mm Hg or on treatment for hypertension.

TABLE 5—Adjusted Odds Ratios for Risk Factors among Asian-American Ethnic Groups: United States Birthplace vs Birthplace in China, Japan, the Philippines, and Other Asian Countries

	Ethnicity							
	Chinese OR		Japanese OR		Filipino OR		Other Asian OR	
Risk Factor	M	F	M	F	M	F	M	F
Body mass index ≥ 24.4 kg/m²	1.6***	1.2	3.0*	1.3	2.0***	1.2	2.4*	2.5*
Hypertension[a]	1.4*	0.9	1.0	0.7	0.9	0.6	3.4*	1.5
Cholesterol ≥ 6.21 mmol/L	1.1	1.0	1.3	0.9	0.8	0.8	1.0	0.4
Glucose ≥ 6.11 mmol/L	1.0	0.9	1.0	1.3	0.9	0.7	1.2	1.3
Smoke any amount (current)	0.6***	2.4*	0.5**	0.7	0.7	1.8**	0.6	1.5
Smoke ≥ 1 pack/day	1.1	4.9**	0.3	1.0	1.0	3.1**	1.9	7.4*

Note. OR = odds ratio. Hawaii was not included in the US birthplace category; see text for discussion of persons born in Hawaii, Hong Kong, Taiwan, and elsewhere. All ORs were computed from coefficients estimated by logistic regression for each ethnicity and gender separately, controlled for age, marital status, education, alcohol intake, and body mass index (except when body mass index ≥ 24.4 kg/m² was the dependent variable).

[a] ≥ 140/90 mm Hg or on treatment for hypertension.

*$P < .05$.

**$P < .01$.

***$P < .001$.

group are more likely to practice US lifestyle habits will be the basis for further discussion.

If the observed smoking differences represent cultural influences, ethnic group differences in smoking behavior probably reflect differences in the respective countries of origin. Asian-American men born in the United States are apparently subject to the same antismoking influences that have resulted in a substantial reduction in smoking among US men of other ethnicities.[14] In a study of ethnic Japanese men,[4,15] a much larger proportions of men in Japan than in Hawaii or California were light smokers, but the proportions of heavier smokers were similar in Japan and California (higher in Hawaii). We found both lighter and heavier smoking to be less prevalent in Japanese men born in the continental United States and Hawaii. Although the two studies used different methodologies, the data suggest that acculturation has resulted in less smoking among Japanese-American men over the 10 to 15 years between the dates of collection of data. We know of no comparable data about smoking that compare US-born and Asia-born persons of other ethnicities.

Why, then, do US-born Asian-American women (except the Japanese) exhibit so much more smoking than their Asia-born counterparts? Some of this can probably be explained by the fact that, among the Asia-born people, the women were much less likely to smoke than the men (Japanese women again excepted). This reflects the situation in the countries of origin.[16,17] It is plausible to hypothesize that the more liberal attitude of American society toward behavior or women tends to equalize this gender disparity in smoking, especially in persons of high educational attainment. Asian-American women born in the United States, under less intense specific social pressures against women not to smoke, apparently take up the habit. In this respect, Asian Americans become similar to other US ethnic groups; among young US White and Black persons, smoking is now almost equally prevalent in the sexes.[14] The Japanese Americans in our study were older and were the only ethnic group with a preponderance of US-born persons; these factors probably account for the smaller gender difference in smoking behavior. Whatever the explanation, these data about smoking suggest an area of major public health concern. The attributable risk of smoking to coronary heart disease is high in young women.[18] By its nature, smoking is the coronary disease risk factor most amenable to complete control. Thus, smoking cessation programs should be targeted toward Asian-American women.

The finding of greater risk of adiposity among US-born and Hawaii-born persons (vs counterparts born in Asian countries) was expected, both on the basis of prior data in Japanese men[4,15] and presumed dietary habits. We did not, however, expect greater adiposity in US-born persons to be limited to men. College education and use of alcoholic beverages were inversely related to adiposity in women but not men in this study. These relations support the argument that acculturation operates in the direction of more emphasis on slimness and fitness in the United States among women but not among men. We have no other data to support this speculation. We do not know whether the interethnic differences we found in the United States reflect similar differences in Asian populations, but suspect they do. Our findings confirm the need to control for body mass index in studies comparing cardiovascular risk factors among ethnic groups.

The only two relevant studies known to us[11,12] agree that Filipino Americans are at greater risk of hypertension than other Asian-American groups. Whatever the explanation, the data have potential implications for research about the pathogenesis of hypertension,[12] as well as pointing to an area for public health efforts.

There are almost no previously reported data concerning cholesterol levels in Asian-American groups. Mean total cholesterol and low-density lipoprotein (LDL) cholesterol levels are lower in Asian countries[19-26] than in Western societies. Higher mean cholesterol levels among ethnic Japanese men living in Hawaii and California (vs those living in Japan) are believed to be due to the higher proportion of fat in the Western diet.[27] We did not find that, independent of body mass index differences, US-born persons of each ethnicity had higher cholesterol levels than their counterparts born in the countries of origin, suggesting either that immigrants from Asia rapidly adopt a high-fat diet or that their progeny maintain similar dietary habits. In support of the former explanation are the similarities of mean total cholesterol levels in US Blacks, Whites, and Asians and scattered reports of rapidly rising fat intake and cholesterol levels in several Asian countries.[22-29]

In this young study population with a majority of women, it will be several more years before the study will have the power to discern possible differences in major cardiovascular disease outcomes. On the basis of the risk factor data herein described. Chinese Americans should fare the best and Filipino Americans and Japanese Americans should fare the worst. For the present, these data can be of use to workers in preventive cardiology. Special efforts should be focused on weight control in Asian-American men, on cigarette smoking among Asian-American women, on diagnosis and treatment of hypertension among Filipino-American men and women, and on control of hypercholesteremia in all Asian Americans. ☐

Acknowledgments

This research was presented in part on November 14, 1988, at the American Heart Association 42nd Annual Scientific Session, Washington, DC; on October 1, 1990, at the 118th Annual Meeting of the American Public Health Association in New York, NY; and on November 12, 1990, at the 44th Annual Scientific Session of the American Heart Association in Dallas, TX. Data concerning blood pressure in Asian Americans have been published.

The authors thank Harald Kipp for computer programming, Cynthia Landy and Rita Coston for data collection, and Lyn Wender for assistance in preparation of the manuscript.

This research was supported by the Alcoholic Beverage Medical Research Foundation Inc, Baltimore, Maryland. The study protocol was approved by the Kaiser Permanente Medical Care Program, Northern California Region, Institutional Review Board.

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Physician Program Cut Infant Death Rate by 45%

Physicians serving poor Miami neighborhoods under a controversial national program cut the infant death rate by 45%, according to a School of Medicine study published in the September issue of *Obstetrics and Gynecology.*

The University of Miami research study is the first documentation of the effectiveness of the National Health Service Corps, a federal program that provides financial assistance to medical students in exchange for a commitment to practice for 1 to 4 years in medically needy communities. The Corps was virtually dismantled by the Reagan administration but has since been partially restored by President Bush in response to pressure from the poor and from physicians.

According to the study, the Corps program saved the lives of at least 320 infants in Miami from 1987 to 1989. Three University of Miami physicians, Hoa Nguyen, Mary Jo O'Sullivan, and Arthur Fournier, examined the impact of placing eight Corps pediatricians and seven obstetricians with four public health clinics: Borinquen Health Care Center in Little Havana, Family Health Center in Liberty City, Community Health of South Dade, and Stanley Myers Community Health Center in Miami Beach.

Before the arrival of Corps physicians, perinatal mortality rates among the poor in Miami varied from 29 to 36 per 1000 live births—about twice the national average of 15.6%. Poor maternal nutrition, financial problems, lack of transportation to clinics, substance abuse and unawareness of pregnancy were cited as contributing factors to the high infant death rate.

Within a year, the mortality rate in the neighborhoods with Corps physicians had fallen to 16 deaths per 1000 live births. In 1988 the rate remained 16 per 1000 and in 1989 it dropped to 13 per 1000, which is better than the national average. There was no significant change in death rates for three other Miami community health centers without Corps obstetricians and pediatricians.

"The Corps physicians were able to follow patients more closely, checking on nutrition and pregnancy complications like diabetes or hypertension. Corps obstetricians were able to supervise deliveries, and pediatricians monitored high-risk babies," said Dr. Arthur Fournier, associate dean for Community Health Affairs at the University of Miami School of Medicine, who served as a Corps physician in rural Onley, Va in 1976 to 1977.

Since the study was conducted, the 15 Corps physicians completed their obligations and left the Miami centers.

"Reagan felt the program was a failure because it didn't keep physicians in the underserved communities after their commitments expired," said Fournier. "But our study demonstrated that for the time they were there, they made a significant impact."

The National Health Service Corps began recruiting volunteers in 1972 to serve in medically needy communities. Scholarship awards peaked during the Carter presidency. But because students did not serve until after they completed their residency training programs, the number of active Corps physicians did not hit its highest point until 1986, when 1400 served. Awards dwindled to a low of 50 in 1988 before rebounding last year. As of May 1991 there are 131 scholarship holders, according to Lynn Trible, public affairs specialist for the Health Resources and Services Administration in Washington, DC. Altogether, 13 804 medical students have gone through the Corps program.

Ethnic Populations in Public Mental Health: Services Choice and Level of Use

Teh-wei Hu, PhD, Lonnie R. Snowden, PhD, Jeanette M. Jerrell, PhD, and Tuan D. Nguyen, PhD

ABSTRACT

Background. Barriers to access and use of mental health care by Asians, Blacks, and Hispanic Americans have been a source of concern for many years. Limitations in our knowledge base persist regarding patterns of use in public sector programs of certain services. Using a sample of almost 27 000 persons, this study examined access and level of use by ethnic minority groups of emergency services, inpatient care, individual outpatient visit, and case management.

Methods. Data from the management information systems of San Francisco and Santa Clara counties were analyzed for fiscal year 1987/1988. Multivariate models were evaluated at two stages, reflecting whether or not a service had been used, and if used, the level of use.

Results. Asians and Hispanics used less emergency and inpatient but more outpatient care than did Whites; Blacks used more emergency and less outpatient care.

Conclusions. Ethnicity continues to play a role in understanding the utilization of mental health services. Regarding emergency and inpatient care, Asian and Hispanic patterns of use appear relatively favorable, whereas the patterns of Blacks continue to be problematic. (*Am J Public Health.* 1991;81:1429–1434)

Introduction

Providing adequate mental health services to the US population has been a national policy for over 20 years. Achieving this objective requires an understanding of factors that affect utilization of mental health services, especially for minority populations. That these populations encounter unusual problems in obtaining mental health services has been acknowledged by the President's Commission on Mental Health.[1] The Commission's 1978 report targeted Asian, Black, and Hispanic Americans for expanded research on the delivery and financing of mental health services and on factors affecting their utilization. More than 10 years later, information on mental health service delivery to ethnic and racial minority populations remains limited. The goal of this article is to provide information on minority utilization of public mental health services that will be useful for planning and policy evaluation.

Previous research on minority utilization of mental health services has focused on several aspects of the problem. Increasingly rigorous studies on community-based samples have tended to emphasize treatment seeking.[2–5] Community-based studies that have gone beyond treatment seeking and looked at patterns of use among those utilizing services have restricted themselves to minority groups in the aggregate[6] or to a single minority group.[7] Also, few studies have focused completely on the public sector.[8]

Large-scale studies of minority utilization of public sector mental health services have been conducted in Seattle, Wash, and in Los Angeles. Sue[9] found that Black, Asian, Native, and Hispanic Americans in Los Angeles County had fewer contacts than Whites with mental health facilities and were more likely not to return after an initial session. In a follow-up of this same county's mental health system 10 years later,[10] O'Sullivan and colleagues discovered major changes. Blacks still failed to return after a single session at a rate higher than Whites, and Asian Americans continued on average to attend fewer treatment sessions after the initial contact. Otherwise, ethnic differences had disappeared. Armstrong and colleagues investigated differences between Blacks and Whites attending a single community mental health center in Seattle.[11] Despite roughly equivalent sociodemographic profiles and similar diagnoses for Blacks and Whites, differences in utilization were found. Whites were more likely than Blacks to have participated in brief therapy, day treatment, activity therapy, and group therapy, as well as in a psychiatric day treatment program and in transitional care. Yamamoto and Silva examined out-

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Teh-wei Hu is with the School of Public Health, Lonnie R. Snowden is with the School of Social Welfare, University of California, Berkeley. Jeanette M. Jerrell is with the Department of Mental Health, Santa Clara County, and Tuan D. Nguyen is with the Department of Public Health, City and County of San Francisco. All authors are associated with the Center for Research on the Organization and Financing of Care for the Severely Mentally Ill, the Western Consortium for Public Health.

Requests for reprints should be sent to Teh-wei Hu, PhD, Professor of Health Economics and Chair, University of California, Berkeley, School of Public Health, Department of Social and Administrative Health Sciences, Berkeley, CA 94720.

This paper was submitted to the journal August 27, 1990, and accepted with revisions July 10, 1991.

TABLE 1—San Francisco County and Santa Clara County Sociodemographic and Mental Health Characteristics of Users of Public Mental Health System by Ethnicity or Race, Fiscal Year 1987–1988

Characteristic	Asian (n = 3825)	Black (n = 3895)	Hispanic (n = 3842)	White (n = 15 146)	Total (n = 26 708)
Gender					
Female	55	48	53	51	52
Male	44	52	47	49	48
Education					
≤ 8 years	46	14	36	25	28
9–11 years	12	25	21	15	17
12 years	17	34	21	24	24
> 12 years	22	23	14	29	25
Mean	8.26	10.25	8.87	10.70	9.52
Marital status					
Married	38	12	25	13	18
Never married	37	62	47	56	52
Wid, Div, Sep	26	27	28	31	29
Age					
18–24	12	16	20	14	15
25–44	51	63	56	57	57
45–64	28	16	18	19	20
≥ 65	8	6	5	10	8
Mean	41.28	36.90	36.42	39.89	38.62
Non-English Speaking	63	2	34	4	16
Selected DSM-III classification					
Schizophrenia	22	30	16	21	22
Affective disorder	18	15	15	20	18
Neuroses	16	8	13	13	13
Adjustment disorder	30	21	32	24	25
Other[a]	15	26	23	22	23
Mean Global Assessment Scale rating (SD)	48.27 (14.89)	43.32 (16.28)	46.67 (17.25)	44.87 (13.32)	45.40 (15.21)

Note. Numbers are percentages. SD = standard deviation. Wid = widowed; Div = divorced; Sep = separated.

[a] Other includes dementia and delirium, alcohol or drug problems, other psychosis, personality, childhood, and deferred.

patient services delivered by the Los Angeles County Mental Health Department.[12] Hispanics received fewer mental health services than Blacks and both groups received fewer services than Whites.

Previous studies of ethnic differences in the use of public mental health systems have several limitations. In particular, few have controlled for group-related differences in diagnosis and level of functioning. Parameters have been estimated without regard for the limitations of dependent variables, including the highly skewed distributions that characterize indicators of mental health service utilization. The problem of utilization has rarely been recognized as consisting of two components: initial use vs. nonuse of a service by clients in the system and level of use among clients with initial use. This article focuses on patterns of public mental health service utilization among Asian, Black, and Hispanic Americans in California's San Francisco Bay area (San Francisco City and County and Santa Clara County). The White population was included in the study as a comparison group for determining distinctive patterns of utilization, after controlling for other potential demographic and clinical differences among the ethnic and racial groups.

Because not every individual used all types of mental health services, this article first studies the probability of using a particular mental health service, and then looks at the amount of use once an individual is engaged in the service. This estimation approach will enable us to address the issue of accessibility of a particular service for a particular ethnic or racial group. We have chosen four major mental health services to study: emergency, inpatient, case management, and individual outpatient services. It is hypothesized that individuals from ethnic and racial minority groups have different probabilities of using these services because of both their ethnic backgrounds and the service policies of the providers.

Methods

Data

This study used two data sources: San Francisco City and County and Santa Clara County for the fiscal year July 1987 to June 1988. Both areas used the Mental Health Billing Information System, which is operated under an interactive data entry and retrieval system. These data are received by clinicians on an episodic basis. The data on amount of use should be accurate because this information is used for billing purposes. However, the *Diagnostic and Statistical Manual of Mental Disorders* (DSM) diagnosis code and the Global Assessment Scale (GAS) may be subject to the clinician's judgment or billing requirements. The reliability and validity of these codings have been reviewed by clinicians at the University of California at San Francisco and reviewed internally by the director of information systems in both counties. The DSM-III[13] ratings were found to be more accurate than the GAS ratings because DSM is the criterion for reimbursements.

Research files were created from raw data files with multiple checks for completeness and accuracy. Both counties have high percentages of ethnic minority populations. Santa Clara has a high proportion of Hispanics, while San Francisco has a high proportion of Asian Americans. Combining the two counties increases the sample size and provides more precision in making empirical estimates. The study excluded persons less than 18 years old and individuals whose treatment was mandated or in any way connected with the criminal justice system. After deleting these observations, the study included 26 708 individuals that used different types of public mental health services. Fourteen percent (n = 3825) are Asian, 15% (n = 3895) Black, 14% (n = 3842) Hispanic, and 57% (n = 15146) White.

Sample Description

Table 1 presents the sociodemographic and mental health characteristics of the sample by ethnicity or race. The sample contains relatively high percentages of Asian and Hispanic females. Asians in the sample had lower educational levels than the other groups. Possibly many Asian immigrants who came to the two counties were older and had lacked educational opportunities in their country of origin. The Asian population had the highest percentage of non-English-speaking users, followed by the Hispanic group.

DSM-III classification[13] and GAS scores are recorded by clinicians during the client's first visit in the 1987–1988 fiscal year. They provide information on initial psychological status. GAS scores ranged from 0 to 99, with higher scores indicating better functional status. The mean GAS score was around 45 and scores were not statistically different among populations. Schizophrenia, affective disorder, and adjustment reactions, the three most common classifications, represent about two thirds of all diagnoses. Asians and Hispanics were more often diagnosed as having adjustment difficulties, perhaps reflecting their immigrant and refugee status, and Blacks were diagnosed more often as schizophrenic. This finding has been reported by Jones and Gray.[14]

Public mental health services include emergency, inpatient, skilled nursing, residential, case management, individual outpatient, medication, and group outpatient services. Of the four services addressed in this article, emergency service is often considered the least desirable type of mental health service; inpatient is the most expensive service; case management is the most important form of community mental health service; and individual outpatient is the most used service.

Table 2 provides information on the percentage of use for these services among the racial and ethnic populations studied. The data show that a higher percentage of Hispanics and Whites (4%) than Asians and Blacks (2%) used inpatient services. A higher percentage of Blacks than other groups used case management (28%) and emergency services (48%). More Asians (66%) and Hispanics (52%) used individual outpatient services than Whites and Blacks. It is interesting to note that among the users of each type of service, the difference in the amount of use among the racial and ethnic groups is not statistically significant.

Data Analysis

The basic data description suggests that differences in the use of the four types of public mental health services among the four populations could be in the percentage of initial use, because there are no significant differences among the users once they are engaged in the service. This preliminary information suggests that one should analyze the data in two stages.

TABLE 2—Percent and Patterns of Use among Users by Type of Public Mental Health Services in San Francisco County and Santa Clara County by Ethnicity or Race, Fiscal Year 1987–1988

Service	Asian (n = 3825)	Black (n = 3895)	Hispanic (n = 3842)	White (n = 15146)	Total (n = 26708)
Percent of use[a]					
Inpatient	2 (70)	2 (93)	4 (140)	4 (531)	3 (834)
Case management	20 (769)	28 (1087)	22 (829)	24 (3564)	23 (6249)
Individual outpatient	66 (2540)	37 (1448)	52 (1997)	45 (6876)	48 (12 861)
Emergency services	22 (858)	48 (1860)	34 (1324)	38 (5727)	37 (9769)
Amount of use[b]					
Inpatient days	12.64 (13.41)	10.37 (11.80)	13.07 (14.21)	12.00 (12.89)	12.05 (13.05)
Case management hours	4.92 (13.55)	4.02 (9.25)	4.59 (13.52)	6.50 (20.88)	5.62 (17.65)
Individual outpatient visits	9.71 (8.50)	8.35 (10.11)	9.04 (10.51)	10.76 (12.05)	10.01 (11.02)
Emergency services visits	1.64 (1.36)	1.76 (1.65)	1.49 (1.15)	1.68 (1.60)	1.66 (1.54)

[a] Numbers in parentheses are number of users.
[b] Numbers in parentheses are standard deviations.

The first stage is to test for the possible ethnic differences in initial use of different types of mental health services. As shown in Table 2, not everyone uses all four types of service. Among the public mental health service users, only 37% used emergency services, 3% used inpatient, 23% used case management, and 48% used individual outpatient services. To differentiate between initial users and nonusers, the dependent variable was designated as 1 for the users and 0 for the nonusers. Logistic regression was used to analyze the selection of services issue.

The second stage of the analysis was to test and measure the differences in the amount of service use once an individual was a user of a given type of service. Because there was a wide variation in the amount of usage, the usage variables were expressed in logarithmic form to adjust for skewness in the distribution. We used least-squares analysis to estimate the amount-of-use equations.

To estimate the real differences among the four ethnic and racial groups, it was important to account for all differences attributable to sociodemographic factors and mental health status, such as age, education, marital status, diagnosis, and county of residence. Although we do not have adequate data on users' income and health insurance status, public mental health services are paid by public funds and a large majority of the users are low income individuals. Thus, the absence of these two variables would not cause bias for the estimated coefficients.

To measure the possible differences among the ethnic populations in the sample, three dummy variables were introduced: Asian, Black, and Hispanic. White is the deleted group used for comparison. We also introduced a dummy variable representing Santa Clara County to account for possible system differences. Because of high correlations between language and ethnicity, we deleted the language variable in the final model.

Results

Table 3 shows the logit analysis of the full model for the probability of using these four types of mental health services. Table 4 shows the least-squares results from the users' sample for these four types of mental health services. All of the explanatory variables (except for GAS) are in categorical form. The comparison groups consist of individuals with the following characteristics: male, White, married, fewer than 8 years education, unemployed, aged 18 to 24, nondisabled, with DSM-III adjustment disorder, and from San Francisco.

To understand the actual magnitude of the effect of these explanatory variables, including the ethnic and racial differences, we calculated each marginal effect (marginal probability) on the use of public mental health services.[15] Only the statistically significant coefficients were calculated with their marginal effects. These results also appear in Table 3.

TABLE 3—Logit Analysis of the Probability of Using Selected Public Mental Health Services among Users in San Francisco County and Santa Clara County, Fiscal Year 1987–1988

	Emergency		Inpatient		Case Management		Individual Outpatient	
Factors	c	m	c	m	c	m	c	m
Intercept	2.351***		−0.783***		−0.989***		−0.908***	
Education (yrs)								
9–11	0.289***	0.066	0.101		−0.118**	−0.021	−0.256***	−0.064
12	0.199***	0.045	0.090	−0.067			−0.237***	−0.059
≥ 13	−0.011		0.247*	0.008	−0.103*	−0.018	−0.043	
Female	−0.193***	−0.044	0.063		−0.094**	−0.017	0.231***	0.056
Asian	−0.762***	−0.174	−0.480***	−0.016	−0.232***	−0.041	0.918***	0.229
Black	0.081*	0.018	−0.186		−0.213***	−0.038	−0.072*	0.018
Hispanic	−0.161***	−0.037	0.006		0.212***	0.038	0.184***	0.046
Never married	0.235***	0.054	0.317**	0.011	−0.407***	0.072	−0.098**	−0.024
Wid/Div/Sep	0.042		0.367**	0.012	0.328***	0.058	−0.021	
Competitive	−0.524***	−0.120	−1.075***	−0.036	−0.131*	−0.023	0.178***	0.044
Homemaker	−1.194***	−0.273	−1.826**	−0.061	−0.510**	−0.091	0.076	
Rehabilitation	−1.433***	−0.327	−1.347**	−0.045	−0.014		0.812***	0.203
School/job	−0.873***	−0.199	−2.902***	−0.097	0.078		0.579***	0.144
Age (yrs)								
25–44	−0.358***	−0.082	−0.454***	−0.015	0.143***	0.025	0.093**	0.023
45–64	−0.671***	−0.153	−1.089***	−0.036	0.365***	0.065	0.007	
≥ 65	−1.307***	−0.298	−2.120***	−0.071	1.184***	0.210	−0.411***	−0.103
Dementia and delirium	−0.024		0.669*	0.022	0.334***	0.059	−0.031	
Alcohol or drug problem	1.373***	0.313	0.565**	0.019	−0.434**	−0.077	−0.828***	−0.207
Schizophrenia	−0.559***	−0.128	0.948***	0.032	1.761***	0.313	−0.151***	−0.038
Affective disorder	−0.373***	−0.085	1.239***	0.041	0.934***	0.166	0.169***	0.042
Psychosis	0.168*	0.038	1.075***	0.036	0.491***	0.087	−0.453***	−0.113
Neuroses	−1.403***	−0.320	−0.239		0.298***	0.053	0.733***	0.183
Personality	−1.149***	−0.260	−0.021		0.492***	0.087	0.590***	0.147
Childhood	−2.403***	−0.549	−0.920*	−0.031	0.170		0.127	
Other	−0.501***	−0.114	−2.092***	−0.070	−0.256***	−0.045	−0.282***	−0.070
Deferred	−2.218***	−0.505	−3.139***	−0.105	0.477***	0.085	−0.534***	−0.133
Global Assessment Scale	−0.029***	−0.007	−0.094***	−0.003	−0.015***	−0.003	0.012***	0.003
Disability (physical)	−0.582***	−0.133	−0.708***	−0.024	0.550***	0.098	0.201***	0.050
Santa Clara resident	−1.299***	−0.297	0.864***	0.029	−1.675***	−0.297	0.335***	0.084
χ^2	5198.07***		2161.84***		4464.20***		2191.66***	

Note. c = logit coefficient; m = marginal effect (marginal probability) where $m_i = c_i(\bar{y})(1 - \bar{y})$ and $\bar{y}$ = sample proportion of the dependent variable. Wid = widowed; Div = divorced; Sep = separated.
*P = .10 level, two-tailed test.
**P = .05 level, two-tailed test.
***P = .01 level, two-tailed test.

Results are discussed by the type of mental health service presented in both Tables 3 and 4, which include the probability of initial use by ethnic or racial differences and the amount of use.

Emergency Services

Table 3 indicates that Asians and Hispanics had less probability (17% and 4%, respectively) of using emergency services than Whites, while Blacks had a slightly higher probability of using emergency services than Whites. Individuals residing in Santa Clara County were 30% less likely to use emergency services than individuals in San Francisco. As shown in Table 4, after the initial use of emergency services, there were no statistically significant differences in amount of use among Asians, Blacks, and Whites. However, Hispanics used 4% less services than Whites. Again, among initial users of emergency services, Santa Clara had 11% fewer users than San Francisco.

Other results in the model indicate that males who were never married, had some high school education (9 to 12 years), were aged 18 to 24, and were diagnosed with alcohol or drug related problems had a higher probability of using emergency crisis intervention. Those with higher GAS scores were less likely to use emergency services. Among the users of emergency services, fewer of the above variables were significant with regard to amount of use, except that individuals diagnosed with schizophrenia and affective disorders used greater amounts of emergency service.

Inpatient Services

The results indicate that Asians had a lower probability (2%) of using public inpatient services. There are no statistically significant differences among

Blacks, Hispanics, and Whites in terms of initial use of inpatient services. Individuals in Santa Clara County had a higher probability of using inpatient services than those in San Francisco. Among inpatient service users, there were no statistically significant differences on amount of use among the four ethnic or racial groups, although patients in Santa Clara County had a 25% shorter length of stay as inpatients than those in San Francisco. Other variables in the model indicate that never married or widowed individuals had a higher probability of using inpatient services, while employed individuals, individuals in school, and the elderly had a lower probability of using these services. Individuals diagnosed with alcohol or drug problems, schizophrenia, affective disorder, and other psychotic disorders had a higher probability of using inpatient services than individuals with adjustment disorder, the deleted group. The higher the GAS score, the less likely individuals were to be admitted to inpatient care. Individuals with physical disabilities were less likely to use inpatient care than the nondisabled. It should be noted that the small magnitude of difference in the probability of using inpatient services among ethnic and racial groups or counties is due to the small percentage of individuals who used inpatient services.

Case Management Services

Case management has become a popular and important form of community mental health service. Case management helps a client gain access to needed medical, social, and educational services. It involves evaluation, development of a service plan, monitoring of the program and movement within the system, and intervention as needed to ensure the availability of the needed services. The results in Table 3 indicate that Asians and Blacks had less probability of using case management services than Hispanic and White individuals. Individuals in Santa Clara had a lower probability (10%) of using case management services than those in San Francisco. Table 4 indicates no significant difference in amount of use among those initially using case management. Individuals with schizophrenia or affective disorder had very high probabilities (31% and 16%, respectively) of initial use of case management. They also used it more often, as did those with adjustment and other psychotic disorders. The higher the GAS score, the lower the probability of case management use.

TABLE 4—Least-Squares Analysis of the Amount of Use for Selected Public Mental Health Services among Users in San Francisco County and Santa Clara County, Fiscal Year 1987–1988

Factors	Emergency	Inpatient	Case Management	Individual Outpatient
Intercept	1.037***	1.967***	0.887***	1.784***
Education (yrs)				
9–11	0.014	0.178**	−0.001	−0.185***
12	−0.011	0.137*	−0.017	−0.139***
≥ 13	−0.012	−0.088	−0.017	−0.056**
Female	0.005	0.097	0.006	0.121***
Asian	−0.012	−0.007	0.005	0.063***
Black	−0.003	−0.137	−0.046	−0.183***
Hispanic	−0.040***	0.050	0.016	−0.112***
Never married	0.004	0.177*	0.171***	0.006
Wid/Div/Sep	0.002	0.170	0.179***	−0.021
Competitive	−0.098***	−0.215	−0.095*	0.151***
Homemaker	−0.071	−0.620	−0.144	−0.071
Rehabilitation	0.063	1.060**	0.074	0.017
School/job	−0.026	−0.268	−0.103	0.159***
Age (yrs)				
25–44	0.001	−0.079	0.066	0.054**
45–64	−0.023	−0.035	0.043	0.062*
≥ 65	−0.059**	0.530**	0.051	−0.065
Dementia and delirium	0.026	0.618**	0.214***	−0.040
Alcohol or drug problem	0.038**	−0.209	−0.001	−0.112*
Schizophrenia	0.170***	0.520***	0.393***	0.197***
Affective disorder	0.116***	0.397***	0.242***	0.146***
Psychosis	0.040**	0.237**	0.157**	−0.126**
Neuroses	0.060***	0.364***	0.124**	0.295***
Personality	0.100**	0.045	0.091	0.270***
Childhood	0.032	0.181	−0.035	0.125**
Other	−0.022	0.443	−0.043	0.031
Deferred	0.059	−0.031	0.227**	−0.003
Global Assessment Scale	−0.003***	−0.004**	−0.001	0.001
Disability (physical)	0.030*	0.138	−0.058	0.009
Santa Clara resident	−0.108***	−0.246***	−0.462	0.001
F-ratio	29.932***	3.268***	13.009***	16.875***
Adjusted R^2	0.1025	0.0836	0.0675	0.0404

Note. Wid = widowed; Div = divorced; Sep = separated.
*P = .10, two-tailed test.
**P = .05, two-tailed test.
***P = .01, two-tailed test.

Individual Outpatient Services

About 50% of the two counties' public mental health clients used individual outpatient mental health services. Among the ethnic and racial populations, Blacks had a lower probability (2%) of using outpatient care than Whites. Asians and Hispanics, on the other hand, were more likely to use individual outpatient services than Whites. Individuals in Santa Clara had a higher probability (8%) of using individual outpatient treatment than those in San Francisco. As to amount of service used, Blacks and Hispanics used 18% and 11%, respectively, less service than Whites, while Asians used 6% more of these services than Whites (see Table 4).

Younger groups had a higher probability of using individual outpatient services than older persons. Individuals with alcohol or drug problems, schizophrenia, and other psychoses had a markedly lower probability of using individual services, while those with neuroses or personality disorders had a higher probability of using individual services. The higher the GAS scores, the lower the probability of using individual services.

Discussion

The findings of this study confirm that there are differences among racial and ethnic groups in choosing public mental health services. Asians had a lower probability than Whites of using emergency, inpatient, and case management services, and a higher probability of using individual outpatient services. Blacks were less likely than Whites to use case management and individual outpatient services, but were more likely to use emergency services. Hispanics had a higher probability of using case management and a lower probability of using emergency services, but there was no significant difference between Whites and Hispanics in use of inpatient and individual outpatient services.

Although there were significant differences in the selection of services among ethnic and racial minorities, there were no significant differences in the amount of either inpatient or case management use. However, in individual outpatient services the differences among ethnic groups continued to exist. Asians used more individual outpatient services, while Blacks and Hispanics used less than Whites. Hispanics used less emergency service than Whites.

Results from the present study provide cause for continuing concern about minority utilization. Utilization by Blacks appears especially problematic, because this group appears to rely more on emergency service and less on case management and outpatient care than Whites. Hispanics in psychotherapy attended fewer sessions than Whites, but in other respects Hispanic utilization patterns were more encouraging. Asians have chosen and used services in ways that were at least as favorable as Whites, and, in most instances, more favorable. Researchers and service planners should avoid generalizing utilization patterns from one minority group to another.

There is a distinct difference between individuals in San Francisco and Santa Clara County. Individuals in Santa Clara County had a lower probability of using emergency services and case management and a higher probability of using inpatient and individual outpatient services than those in San Francisco. Santa Clara County individuals used lesser amounts of emergency, inpatient, and case management services than users in San Francisco. The differences between areas could be explained by the decentralized nature of providing public mental health services in California. The flexibility in the state's funding policy allows each county to deliver different levels and types of mental health services. In San Francisco, the city allocates very limited resources to inpatient services and places a greater emphasis on the broad use of case management as an appropriate service modality. On the other hand, Santa Clara County provides more generous inpatient services and has created eight new community mental health outpatient service centers for Asian and Hispanic minorities. Consequently, patients in public health systems in San Francisco and Santa Clara counties receive different types and amounts of mental health services. □

Acknowledgments

This research was supported by grant no. R01-MH44753 from the National Institute of Mental Health. We wish to thank C.Y. Liu for his computer programming and Jacqueline M. Golding, Richard M. Scheffler, and Paul Widem for their helpful comments.

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Income, Race, and Surgery in Maryland

Alan M. Gittelsohn, PhD, Jane Halpern, MD, and Ricardo L. Sanchez, MD

ABSTRACT

Background. We describe common surgical and medical hospital admission rates for Maryland residents, exploring systematic effects of race and income.

Methods. The data comprise Maryland hospital discharges and population estimates for 1985 to 1987. Patient income is the race-specific median family income of residence zip code. Logistic regression is used to measure incidence by race, income, and residence for surgical and medical reasons for admission.

Results. Population rates for discretionary orthopedic, vascular, and laryngologic surgery tend to increase with community income levels. Coronary and carotid artery surgery rates are two to three times higher among Whites. The more discretionary the procedure, the lower is the relative incidence among Blacks. By contrast, admission rates for most medical reasons decline with increasing income levels and are elevated among Blacks. The affluent receive coronary artery procedures whereas the poor are hospitalized for coronary artery disease.

Conclusions. Blacks and the poor appear to have higher illness burdens requiring hospital care. Discretionary surgeries have a White predominance and increase with income; medical admissions have a Black predominance and decline with income. Race and community income level are important factors in differential hospital utilization rates. (*Am J Public Health.* 1991;81:1435–1441)

Introduction

Significant variations in the incidence of common surgical procedures have been observed in the United States, Canada, and Western Europe over the past 50 years, a general finding being intermediate to high variability in surgical rates between population groups.[1–9] The predominant explanation for geographic differences has emphasized physician practice styles. Populations may differ in disease prevalence, demographic characteristics, access to care, and medical resources. These factors are confounded because measurement in geographic studies involves place of residence or exposure. The result is that a high or low surgical rate for a community cannot be ascribed directly to morbidity, access, or medical supply variables. High-rate areas have been suggested as representing "unnecessary" surgery and low rate areas as "underserviced." Underservice has been proposed to explain the equal distribution of "appropriateness" of care observed for Medicare patients receiving procedures in high- and low-rate communities.[9]

Unresolved questions pertain to the interpretation of variations in surgery rates in terms of health benefit, questions that cannot be answered with incidence data alone. The efficacy of some procedures and indications for appropriate intervention have yet to be established. Studies of the variability of surgical opinion report a lack of concordance for several common surgeries.[10] Few procedures have been studied by randomized clinical trials to assess benefits and delineate indications for which benefits might be derived. Without efficacy studies, it is not possible to determine whether a high-rate area indicates overuse or a low-rate area underservice. We include common medical reasons for hospital admission for comparison and as indicators of need for surgery. Our purpose is to describe use of surgical and medical services by race, income, and geographic area in Maryland.

Methods

Data Sources

We derived surgical and medical admission rates from 1985 to 1987 discharge abstracts, collected by the Hospital Services Cost Review Commission (HSCRC), for Maryland patients admitted to acute care hospitals. The abstract includes age, race, sex, zip code of residence, procedures, diagnoses, and diagnosis-related group (DRG). Records for Maryland residents treated in Washington, DC, and northern Virginia hospitals in 1985 are included. These are important for Washington suburbs, where patient migration into the District of Columbia is common, particularly for specialized care provided in the medical centers of the city. Surgical cases are classified by the most important procedure and medical admissions by DRG. We use "incidence" as the count of occurrences of a surgery or DRG. The admission rate is incidence relative to the population at risk by age, sex, race, income, and zip code of residence, irrespective of the hospital providing care.

The authors are with the Departments of Biostatistics and Health Policy, Johns Hopkins School of Hygiene and Public Health, Baltimore.

Requests for reprints should be sent to Alan Gittelsohn, PhD, Department of Biostatistics, Johns Hopkins University School of Hygiene and Public Health, 615 North Wolfe Street, Baltimore, MD 21205.

This paper was submitted to the journal August 13, 1990, and accepted with revisions April 4, 1991.

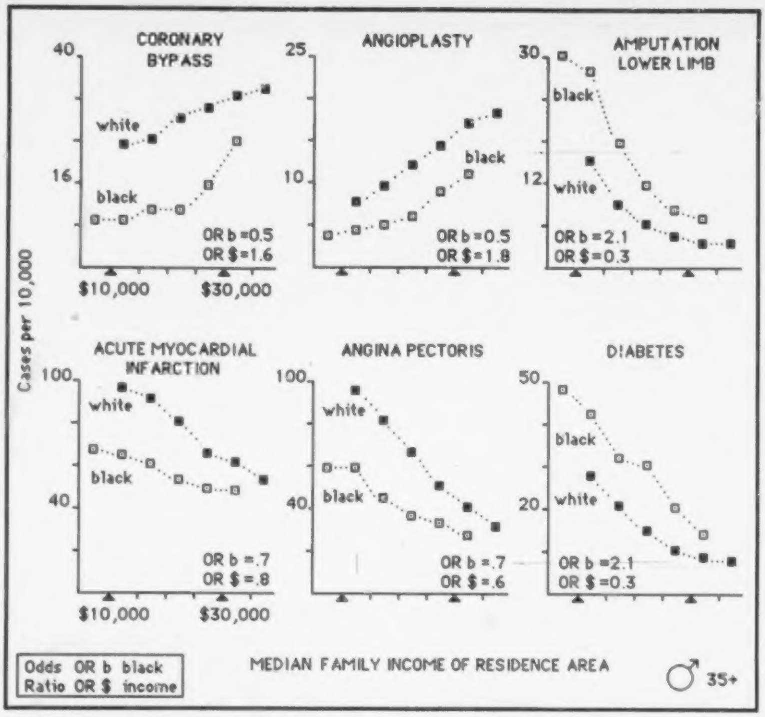

FIGURE 1—Hospital admission rates by race and income, Maryland men aged 35+, 1985–1987.

Population

Maryland's 1986 population of 4.4 million was distributed in 485 zip code areas, grouped into 115 areas by combining adjacent small zip codes. Populations were derived from the 1980 census and 1986 county projections by age, race, and sex.[11] Nonwhite nonblacks constitute under 5% of non-Whites and are classified as Black. Blacks, who make up 27% of Maryland's population, predominate in low-income areas constituting the majority of the poor. Appalachian western counties and high-income suburbs have few Blacks. Four Baltimore City zip codes are 90% White and three are more than 90% Black. We used race-specific median family income of zip code of residence in our analysis because income was not available in the patient record. Zip code median incomes for Whites ranged from $7000 to $58 000, medians for Blacks averaged one third less. We estimated separate medians by race because of wide income gaps between Blacks and Whites living in the same zip code. Variability within areas limits the value of group median income, possibly resulting in attenuation of relationships of surgery and DRG rates with income. There were no ready means for projecting 1979 census incomes to 1986 to account for declines among the poor and increases among the wealthy. Few problems arose in uniformly poor areas and in homogeneous suburbs. Problems did occur in zip codes with wide income variation and pockets of affluence or poverty.

Statistical Methods

The indirect surgical rate for an area is the ratio of observed/expected cases; expectations were based on Maryland age-specific rates. The distribution of 115 area case ratios for a procedure around the state average of 1.0 includes random and systematic components of variance described by MacPherson.[6] The random component depends on the number of cases, whereas the systematic component (SCV) reflects systematic effects above random variation. Incidence is related to multiple factors through coefficient beta in logistic regression.[12] A case is coded "1" and a population datum "0" within age, race, sex, income, and area classes. The estimate b and its standard error SE(b) provide a test of the hypothesis of no effect of the factor. The odds ratio OR = Exp[rb] is relative incidence over the range (r) adjusted by the factors in the model. With $r = 1$, OR(male) compares males with females (M = 1 and F = 0) and OR(Black) compares Blacks with Whites (B = 1 and W = 0). Age is coded in 10-year groups from under 5 years to 75 years and over. For income coded in $1000 units, we used $r = 20$ with an interval of 10 to 30 because there were few Whites living in areas with a median below $10 000 and few Blacks living in areas above $30 000. OR(income) = Exp[20b] compares areas with medians of $30 000 with areas with medians of $10 000.

Race and Income Trends

We use first order models for simplicity of presentation. Small deviations of the betas from 0 and minor differences between slopes are often significant when regression is applied to large populations. Figure 1 shows examples of surgical and DRG admission rates with differing income trends by race. Odds ratios are included for comparison with the trends. The rates pertain to men over 35 years of age, with the 115 zip code areas grouped into $5000 income classes. Coronary bypass and angioplasty rates increase over 60% with income for both races. With an OR of 0.5, Blacks have lower rates within each income class, indicating presence of a race effect over and above income. Nonparallel income trends by race are seen, with race differences that are pronounced among low-income groups and reduced among higher income groups. By contrast, admissions for acute myocardial infarction and angina pectoris decline with income and are reduced among Blacks. Thus, coronary artery disease is high among the poor and coronary artery surgery is high among the more affluent. A different set of relationships is present for diabetes and the complication of limb amputation for metabolic or endocrine reasons. The income OR of 0.3 indicates marked declines in diabetes and amputation rates over incomes and the Black OR of 2.1 indicates twofold higher rates among blacks. The race–income interactions are small but statistically significant. Their introduction into logistic models adds complexity without clarification. The income OR is an average for the races and the Black OR is an average over incomes.

Results

Small Area Variation in Surgery Rates

Figure 2 illustrates geographic distributions with frequency histograms of case ratios for Whites in the 115 Maryland areas. The distributions describe total variation, including random components and SCVs. More than half of the areas have few Blacks, particularly suburban, rural, and western areas. Heterogeneity is reduced by limiting the figures to Whites. We include common surgeries and surgeries with low and high geographic variance. Coronary bypass, cholecystectomy, and hysterectomy are the leading Maryland surgeries based on total hospital charges per capita.

Procedures among Women

Mastectomy (simple and radical) had the lowest variance of the surgeries studied (SCV = .03). The low variance reflects uniform incidence and detection of breast cancer and medical consensus as to its management. Reduction mammoplasty had the highest variance observed (SCV = 1.07), its use increasing markedly with income. Generally, cosmetic procedures have high systematic variances based on patient demand and income. Cesarean section had low geographic dispersion, indicating similar obstetrical approaches throughout the state. Hysterectomy had moderate variance (SCV = .08). Cholecystectomy varied twofold between moderate-income South Baltimore and affluent North Baltimore. Variation may be explained by different social factors, approaches to management, or incidence of cholelithiasis in these populations.

Procedures among Men

Prostatectomy and inguinal hernia repair were nearly uniform across the state. The low SCVs suggest uniform prevalence and apparency of indications, agreement as to management, and few barriers to care. Nondiscretionary appendectomy had an outlier in Cambridge, with eight times the rate of nearby Annapolis. The high surgery rate may be due to high incidence of appendicitis, to a lower threshold for surgery, or to a high rate of false positive indications. Laminectomy was discretionary, with incidence varying fourfold between southwestern Baltimore and the Eastern Shore. The latter region may be underserved by orthopedic surgeons. Differentiating overuse in one area from underuse in another implies study of appropriateness and geographic and occupational variations in spinal disk problems. Coronary bypass had high geographic variability (SCV = .63). Explanation requires distinguishing among disease effects, access, referral, and criteria for case selection.

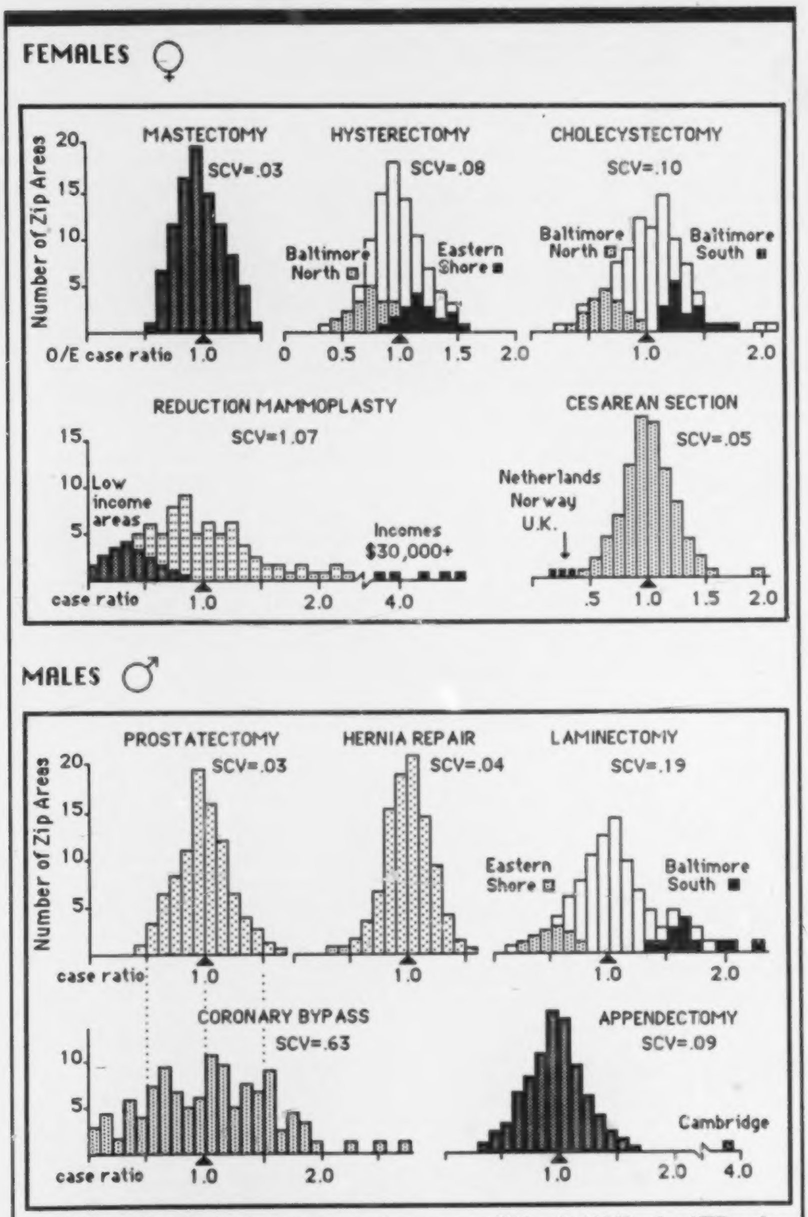

FIGURE 2—**Relative incidence of procedures among Whites in 115 Maryland ZIP code areas, 1985–1987.**

Variability of case ratios among the 115 small areas was observed for each procedure by statistical test, including low-variation mastectomy and inguinal hernia repair.[illegible,13] For common surgeries, small deviations from randomness are significant because of sample size alone. Tests for uniformity of rates provide little new information because the null hypothesis is almost always rejected. Local variations in surgical practice were present for many procedures. The distributions reflect reported services as delivered in Maryland during the study period. The low variance of cesarean section rates (SCV = .05) points out a limitation of small area analysis. The state average is a yardstick for comparing the 115 areas, not an objective standard. The Maryland cesarean rate is two to three times that of Western European countries.[14] We found no definitive results relating obstetrical practice differences to birth outcomes on a population basis. A similar situation held for many discretionary surgeries.

TABLE 1—Relative Incidence of Common Surgeries: Odds Ratios for Race and Median Family Income, Maryland, 1985–1987

Sex	Procedure	No. of Cases	Odds Ratio Income (high/low)	Odds Ratio Black (B/W)
Both	Sinus procedures	7781	1.23*	0.80*
	Total hip replacement	7761	1.22*	0.73*
	Tonsillectomy	22 743	1.13*	0.44*
	Hemorrhoidectomy	12 105	1.05	0.69*
	Appendectomy	13 112	1.01	0.64*
	Gastrectomy	2972	0.68*	1.47*
	Amputation of limb	7077	0.29*	2.03*
Female	Cosmetic surgeries	8135	4.53*	0.30*
	Operations on foot and toe	3655	1.58*	1.97*
	Mastectomy	6665	1.49*	0.96
	Thyroidectomy	2857	1.48*	1.51*
	Hysterectomy	28 891	1.21*	1.42*
	Cholecystectomy	15 776	0.54*	0.56*
Male	Angioplasty	3439	1.82*	0.52*
	Coronary bypass	7209	1.62*	0.45*
	Laminectomy	7441	1.42*	0.49*
	Prostatectomy	17 762	1.12*	0.99
	Inguinal hernia repair	14 628	1.02	0.78*
	Carotid endarterectomy	2244	0.83*	0.31*

Note. Odds ratios are based on logistic regression of surgery by age, race, sex, and income. Cosmetic procedures include blepharoplasty, rhinoplasty, mammoplasty, and facial rhytidectomy.

*$P < .01$

Surgery Rates by Income and Race

Table 1 is a summary of logistic regressions for common surgeries showing ORs for income and race. Sex-dominant procedures such as mastectomy for women and hernia repair for men were computed for one sex. Most surgeries increased or remained constant over income levels. Cosmetic surgery rates exhibited fourfold and higher variation between low- and high-income areas, with access explaining the difference. For vascular procedures, the income OR was 1.8 for angioplasty, 1.6 for coronary bypass, and 1.5 for coronary angiography. The income ORs were 1.6 for sinus procedures and 1.3 for tonsillectomy. Orthopedic procedures involving hip, knee, and spine increased with income. Inguinal hernia repair, appendectomy, and prostatectomy were not related to income. The income OR was 0.5 for cholecystectomy, one of the few discretionary surgeries that declined with income. Women in high-income areas had half the gallbladder surgery rate of women in low-income areas. The lowest surgical income OR was 0.3 for nondiscretionary amputation of limb for endocrine and metabolic reasons. Many of these cases were diabetics with peripheral vascular disease.

Whites had higher rates than Blacks for most discretionary surgeries. The low Black OR for elective surgeries is explained by income or by deselection of Blacks for the procedures. Hernia and appendectomy rates were lower among Blacks. Laryngologic surgery was elevated among Whites. Tonsillectomy rates were 29 per 10 000 for Whites and 13 for Blacks. Orthopedic surgery was more common among Whites. The laminectomy Black OR of 0.5 indicates that half as many procedures were performed among Blacks as among Whites. Adjusted by age, sex, and income, the 1:2 Black to White ratio implies that race was a contributing factor per se. Race differences in vascular surgery were measured by two to three times higher rates among Whites for angioplasty, coronary bypass, and carotid endarterectomy. Prostatectomy and mastectomy exhibited no race difference. Gynecologic surgery was higher among Black women, with a Black OR of 1.2 for hysterectomy. Foot operations, including excision of bunion and arthroplasty, were twice as common among Blacks. Either Blacks had more foot problems requiring care or they received more inpatient surgery because of differing medical practice. Nondiscretionary gastrectomy and limb amputation were elevated among Blacks (ORs = 1.5 and 2.0). Discretionary aspects of common vascular, orthopedic, and laryngologic procedures suggest that lower rates for Blacks were related to referral and access rather than morbidity.

Medical Admissions

Table 2 shows ORs for common medical DRGs. Persons living in low-income areas were more likely to be hospitalized for medical reasons than persons in high-income areas. This pattern suggests an income relationship whereby the poor have a higher incidence of conditions requiring hospital care. For all medical DRGs, the income OR of 0.6 was a decline of 40% over incomes from $10 000 to $30 000 per year. Angina pectoris, hypertension, and heart failure rates decreased by more than half with income. The income OR was 0.3 for pneumonia and 0.2 for bronchitis and asthma. Rates for alcohol-related DRGs declined by 75% with income, in part reflecting use by those privately insured of treatment facilities not in the HSCRC database. Gastrointestinal obstruction, hip fracture, and medical back problems were among the few common medical DRGs that did not decline in incidence with income.

Women predominated for hip fracture, bowel inflammation, urinary tract infection, hypertension, and diabetes. The high female rates for hip fracture occurred among Whites but not Blacks. Men predominated among heart disease (male OR = 2.0 for acute myocardial infarction). Male-to-female ORs above 3.0 were noted for cirrhosis of the liver and alcohol-related DRGs. Although Blacks had lower surgery rates than Whites, the total medical admission rate was 22% higher for Blacks. Blacks had high ORs of 2.0 for diabetes, 3.0 for hypertension, 2.3 for alcohol-related DRGs, and 5.4 for red blood cell disorders, a difference due to sickle cell anemia. An exception was hip fracture (Black OR = 0.4 for women and 0.8 for men). Among cardiovascular conditions, Whites had higher rates for acute myocardial infarction, angina pectoris, and arrhythmias and Blacks had higher rates for heart failure.

Discussion

Explanations for geographic variations in surgical and medical admission rates have emphasized the role of physician practice style differences reflecting medical uncertainty regarding optimal treatment regimens. Our results indicate that race and income may confound differences between small areas, particularly in states with heterogeneous populations. These demographic factors may influence medical need in terms of disease prevalence and severity as well as the demand

TABLE 2—Relative Incidence of Diagnosis-Related Groups (DRGs): Odds Ratios for Race, Sex, and Median Family Income, Maryland, 1985–1987

		Odds Ratio		
Diagnosis-Related Group/Medical DRG	No. of Cases	Income (high/low)	Male (M/F)	Black (B/W)
Nervous system	80 421	0.6*	1.1*	1.0
Transient ischemic attack	10 842	0.8*	1.1*	0.9*
Seizure and headache	16 082	0.4*	1.0	1.0
Circulatory system	178 270	0.7*	1.4*	0.9*
Acute myocardial infarction	32 297	0.8*	2.0*	0.7*
Arrhythmia	20 306	0.6*	1.3*	0.9*
Angina pectoris	35 737	0.5*	1.1*	0.8*
Hypertension	5908	0.5*	0.7*	2.8*
Heart failure and shock	33 820	0.4*	1.2*	1.2*
Gastrointestinal system	110 086	0.7*	1.0	1.0
Obstruction	6721	1.1*	0.8*	1.1
Hemorrhage	13 078	0.8*	1.5*	1.3*
Cirrhosis of liver	2756	0.2*	3.1*	1.2*
Respiratory system	122 385	0.4*	1.2*	1.0
Pulmonary edema	17 928	0.4*	1.6*	1.0
Pneumonia and pleurisy	32 334	0.3*	1.2*	1.0
Otitis media and URI	8053	0.3*	1.0	0.5*
Bronchitis and asthma	38 451	0.2*	0.9*	1.0
Other medical DRGs				
Hip fracture	2791	1.2*	0.6*	0.6*
Medical back problems	40 659	1.0	1.1*	0.8*
Red blood cell disorders	9422	0.8*	1.0	5.4*
Urinary tract infections	17 597	0.5*	0.4*	1.0
Diabetes	17 343	0.3*	0.8*	2.1*
Alcohol-related DRGs	19 446	0.3*	3.6*	2.3*
All medical admissions	1 478 027	.6*	1.2*	1.2*

*$P < .01$

for and access to primary and secondary care. The surgery rate for a community is based on multiple factors, including demography, morbidity, access to care, referral practice, physician decisions concerning patient management, and medical resources.

Generally, small area data are unavailable for incidence or prevalence of conditions leading to hospital care. An exception is cancer measured by mortality statistics and tumor registries. Objective definition of disease by pathology, high levels of ascertainment, and concordance among physicians concerning management lead to correspondence between surgical rates and breast cancer incidence. Race differences in mastectomy rates may be related to incidence of breast cancer by income operating through differences in fertility. A review of the role of pregnancy history reports protective effects of early pregnancy and high risk among nulliperous women.[15] Deferred child bearing and lower fertility among the more affluent may explain the income OR of 1.5 for mastectomy, one of the few nondiscretionary surgeries increasing with income. Morbidity explains the distribution of mastectomy. By contrast, there is little information on the occurrence of cholelithiasis, prostate hypertrophy, uterine fibroids, and other indications for common surgeries. Morbidity cannot be ruled out as a determinant of surgical rate differences between communities. Uncertainty holds for the relative contributions of disease factors, selection of cases for particular types of care, access, referral networks, and physician practice styles.

Access to care involves patient perceptions, financial means to secure care, availability of care, and the referral network to surgeons, which, if constrained, may result in low surgery rates. Such constraint may hold for Maryland Blacks for elective, vascular, orthopedic, and laryngologic surgery. Most surgeries exhibit either a positive relationship or no relationship to income. An implication is that access factors related to income are determinants of use of discretionary procedures. Among common discretionary surgeries, cholecystectomy has the largest decline with income. The decline for Whites is consistent with a disease hypothesis whereby residents of high-income areas have reduced incidence of cholelithiasis. Management of gallbladder disease and referral behavior in low-rate areas may restrict surgical intervention. Residents of affluent communities are not constrained in access because of limitations in income and insurance coverage. Cholelithiasis increases with obesity and is reduced among higher social classes.[16] Among women, Whites have 50% higher cholecystectomy rates than Blacks despite lower incomes and higher obesity levels among Blacks.[17] Race differences in cholecystectomy may be due to reduced occurrence of gallstones among Blacks or to deselection of Blacks for surgery. Blacks have lower rates for elective, vascular, orthopedic, and laryngologic procedures, whereas parity is seen for prostatectomy and a 22% excess for hysterectomy. The more discretionary the procedure, the lower the relative incidence among Blacks. Discretionary aspects of vascular, orthopedic, and laryngologic procedures and lack of evidence supporting morbidity differences by race suggest that the low rates for Blacks relate to referral and access. There are no definitive studies documenting race differences in prevalence of lesions of coronary and carotid arteries. Mortality data indicate that Blacks have similar or higher death rates than Whites for ischemic heart and cerebrovascular disease. Access, referral practices, and case selection may explain the two to three times higher incidence of vascular surgery among Whites.

Rates for most medical DRGs decline with income, a finding consistent with higher morbidity among low-income groups. Diabetes, bronchitis, and asthma and pneumonia rates decline by two thirds with income. Since diabetes and asthma are thought not to depend on social class, medical management may underlie the declining hospital rates with increasing income levels. In contrast to lower Black rates for most surgical specialties, Blacks have higher rates than Whites for many common DRGs. The relationship between illness measured by hospital admission and surgical use is variable. There is a twofold higher incidence among Blacks for diabetes and the complication of limb amputation for endocrine and metabolic reasons. Coronary bypass, angioplasty, and angiocardiography have similar patterns of use, increasing with income and significantly elevated among Whites within each income group. Race effects are present over and above income. Coronary artery surgery (CAS) increases with income, whereas coronary artery disease (CAD) declines. High CAD rates occur in low-income areas and high CAS rates in high-income areas, a discordancy between treatment and need as measured by these means. Blacks have 20% lower rates for CAD and markedly lower rates for CAS (Black OR = 0.4 for coronary bypass and 0.5 for angioplasty). The race difference for CAS is highest among low-income groups and approaches parity among the more affluent. Maryland Blacks have access to hospital care for medical reasons and are deselected for discretionary procedures in at least three surgical specialties. Morbidity explanations for most of the medical and surgical admission rate differences by race and income remain to be elaborated. A general conclusion is that the medical care system of Maryland constrains the use of discretionary vascular, orthopedic, and laryngologic surgery by the poor and by Blacks.

Incidence rates based on hospital abstracts and population projections have several limitations. Data problems are inherent in an administrative health data system with 55 participating hospitals. A major constraint is inability to link records for multiple admissions of a patient. Patients seeking care in other states are not reported. A check of border areas does not indicate abnormal rates. No information is available on quality of abstracting. Formal errors are present with impossible combinations of ICD codes by sex and age (e.g., men with gynecologic diagnoses and the aged with perinatal conditions). Errors of this type are rare compared with factual errors undetected by editing. Experience in other settings may be relevant. After the swine flu affair, record reviews of abstracts with Guillain-Barre syndrome and abstracts with tuberculosis as comorbidity revealed nearly 100% false positives due to miscoding.[18] A reliability study comparing records with independent reabstracts found three-digit ICD agreement ranging from 34% for chronic ischemic heart disease to more than 90% for surgical diagnoses.[19] It is likely that error rates are now reduced through more careful scrutiny and use of records in reimbursement. There are uncertainties regarding populations and assignment of residence. There are abstracts with zip codes not in the postal directory and no zip codes for the homeless. Baltimore's Central Post Office is zip 21203 coded for patients by HSCRC but with no census count and over 5000 PO boxes, the size of a town. Questions pertain to 1986 county projections applied to small areas because population growth and loss rates are not uniform for zip code areas within Baltimore City or suburban counties. Use of zip code median income for individual patient income is a major, unavoidable limitation. As unassembled intelligence, the data require caution in drawing conclusions. For these reasons, we emphasize large differences not likely to be explained by artifacts of measurement.

We used population rates because of epidemiological interest and concern regarding selection bias with inpatient data. Population rates are resultants of morbidity, access, and selection factors, whereas hospital rates relate to preselected inpatients. Case fatality depends on treatment and prognosis, with clinical details on severity not routinely available. Two recent inpatient studies illustrate ambiguity due to possible selection bias. A case fatality study for uninsured patients assigns prognosis using diagnosis, age, race, sex, and comorbidity.[20] Higher death rates are reported for the uninsured. The authors note that the mortality difference "also could be due to differences in severity of illness between the uninsured and privately insured," a reference to selection bias. It is plausible that the uninsured defer seeking care and have less favorable prognoses at time of hospitalization. A CAS study explores "whether there are different procedure rates for black patients once they are hospitalized for serious cardiovascular conditions."[21] The series consists of patients with coronary artery procedures or with principal diagnosis of circulatory disease, including ischemic heart disease (IHD) and non-IHD circulatory diseases. Non-IHD patients with hypertension or cerebral hemorrhage ordinarily do not receive CAS. The authors show CAS rates of 0.4% for non-IHD and 39% for chronic ischemia patients, a hundredfold difference. There is an excess of Blacks in the low-risk groups. In Maryland, most CAS cases have coronary ischemia with selection and surgery scheduled before, not after, admission. Thus, $\Pr\{\text{CAS}|\text{coronary ischemia \& CAS scheduled}\}$ is about 1 and $\Pr\{\text{CAS}|\text{hypertension or stroke}\}$ is about 0. We concur with the finding of race differences in use of CAS but based on population studies. Population-based rates avoid some of the uncertainties involved in measuring risks among preselected patients. Central issues in population studies are factors leading to selection for hospitalization.

We have described variations in surgical and hospital care by race and income for Maryland during 1985 through 1987. Our results show that the more discretionary the procedure, the lower its relative incidence among Blacks and the poor. Admissions for common medical reasons are elevated among Blacks and decline with increasing income for both races. Varying rates between communities reflect the interplay of diverse factors, including patient behaviors, morbidity, access, physician practice styles, referral, and resource availability. Each factor has different implications for strategies designed to reduce health care costs, to improve quality of care, and to help ensure equitable allocation of care. Incidence data using hospital records and population estimates alone cannot clearly distinguish between these effects and can only provide indicators of unexplained variability and measures of low and high occurrence rates. Incidence data can be used to target potential problems, to assist in assigning priorities for special investigation, and to provide descriptions of medical care impacts on population groups. ☐

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Effects of Budgeting on Health Care Services in Dutch Hospitals

Anton F. Casparie, MD, PhD, and Dick Hoogendoorn, MD, PhD

ABSTRACT

Background. In 1983 hospital budgeting was introduced in the Netherlands. We studied the effect of the enactment of budgeting on the efficiency and effectiveness of health care.

Methods. In four different age groups, the admission rate, length of stay, type and number of surgical inpatient procedures, and hospital mortality were measured in all short-term hospitals from 1977 through 1988. Data were standardized by age and sex.

Results. For the total population, the hospital admission rate and the operation rate decreased after 1982. However, for the subgroup of patients beyond the age of 65 both rates are still on the rise, but the increase in the admission rate for elderly patients has slowed significantly since 1983. The tendency toward a shorter length of stay, together with the diminished admission rates, led to a 22% decrease in standardized hospital days between 1982 and 1988. The severity of the operations increased. Most operations performed on elderly patients were aimed at improving the quality of their lives rather than lengthening their life expectancy. The hospital mortality rate decreased in all age groups.

Conclusions. The findings suggest that modern medicine in the Netherlands has become more efficient and more effective. Better health care for older patients was achieved within the same budget. The tendency toward more efficiency by hospitals has been reinforced since 1983. (*Am J Public Health.* 1991;81:1442–1447).

Introduction

In the Netherlands, a country with nearly 15 million inhabitants, as in many other countries, the government and policymakers have been concerned about the increase in the costs of health care in the past few decades. One effort by the Netherlands to control these costs and to stimulate efficient health care was the introduction of hospital budgeting in 1983. Prior to 1983, hospitals were financed on the basis of their production output, a patient-day price depending on contracts with sickness funds.

In 1983 and 1984, a sum was allotted to each hospital, based only on what it had spent in the previous year on the treatment of both inpatients and outpatients. In this system a hospital received a global budget for all its activities.[1] Since 1985, the budget has been calculated by a combination of a fixed and a variable part. The fixed part was defined by two capacity input parameters: the number of beds and specialist units for clinical and ambulatory care. Output parameters such as number of admissions and patient days determine the variable part of the budget. This variable part was determined in advance in negotiations between hospitals and local sickness funds and depended on the expected number of output parameters. All the activities had to be carried out within this global budget. If the expenses of a certain year exceeded the budget by 5% or more without an evident explanation, the budget was decreased the following year (for more details see Maarse[1]).

The first aim of the present study was to assess how budgeting has affected the admission rate and length of stay and the frequency and type of surgical procedures. As the medical profession is the principal determinant of the use of health-care services, the effects of budgeting will ultimately be reflected by changes in the practice of medical specialists. The second purpose of this investigation was to examine the relation between budgeting and the effectiveness of care as reflected by hospital mortality.

Since 1964, in the Netherlands, the Centre for the Registration of Medical Data from Hospitals has gathered inpatient data on age, sex, diagnosis on discharge, length of stay, type of surgical procedure, and in-hospital death. All general, short-term Dutch hospitals have joined this center on a voluntary basis.

In 1982, there were 176 general short-term hospitals; 90% were private. The hospitals vary in size from 60 to 900 with an average of 300 beds (4.5 beds per 1000 population).[2] About 70% of the population (those under a specified income) is covered by compulsory insurance with the sickness funds. The remainder of the population is insured by private companies. Medical specialists work within a closed staff model: they practice full-time, usually in one hospital. The medical staff meets regularly, usually once per month. Patients have access to hospital outpatient departments staffed by specialists only through referral by general practitioners.

Anton F. Casparie is with the Erasmus University Rotterdam, Department of Health Care Policy and Management. Dick Hoogendoorn, at the time of this study, was with National Centre for the Registration of Medical Data from Hospitals, Utrecht, the Netherlands. Dr. Hoogendoorn has since died.

Requests for reprints should be sent to Prof. dr A. F. Casparie, Department of Health Care Policy and Management, Erasmus University Rotterdam, PO Box 1738, 3000 Dr Rotterdam, The Netherlands.

This paper was submitted to the journal December 20, 1989, and accepted with revisions March 13, 1991.

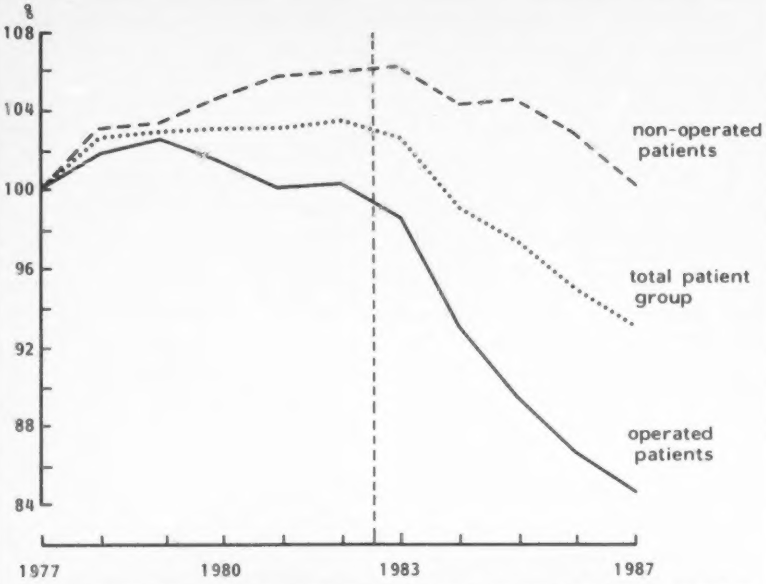

FIGURE 1—1977 age-adjusted admission rates for patients undergoing an operation (51.0/1000 population) and patients not undergoing an operation (56.2/1000 population) divided by comparable rates for each year (1977–1988) and multiplied by 100.

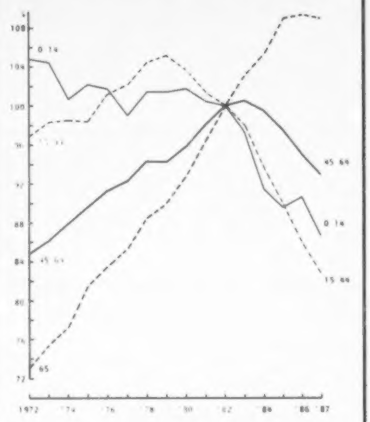

FIGURE 2—1982 admission rates for age groups 0–14 (101.0/1000 population), 15–44 (88.8/1000 population), 45–64 (123.4/1000 population), and 65 years and over (207.1/1000 population) divided by comparable rates for each year (1977–1988) and multiplied by 100.

General practitioners themselves do not have hospital admission privileges. In the Netherlands, most medical specialists work on a fee-for-service basis; for outpatient treatment of publicly insured patients, the specialist receives a fixed amount of money that covers 1 month of treatment. These arrangements remained unchanged after the enactment of hospital budgeting.

Methods

Data regarding the population were obtained from the Dutch Central Bureau of Statistics.[3] The data from the Centre for Registration of Medical Data from Hospitals have been standardized by age and sex and classified by diagnosis according to the international Who-classification, ICD-8 (international classification of diseases, eighth revision) and ICD-g-CM (ninth revision, classification of mortality). All invasive therapeutic procedures, with the exception of infusions, are classified as operations. Seven classes of surgical intervention have been weighted from 1 (easy) to 7 (very demanding), with the average length of time needed for the operation serving as the criterion of relative weight.

Following standardization, annual rates were divided by comparable rates in a given (base) year and multiplied by 100. Thus, the figures reflect percentages with the base year as 100%. The observed numbers of admitted patients and of operated patients per 1000 (y) have been regressed against the time (x) in years, for the periods 1972 through 1982 and 1983 through 1988. The slopes and the 95% confidence intervals for the differences between these slopes were determined with the Confidential Interval Analysis Program of the *British Medical Journal.*

Results

Admission Rate

For the total population, the hospital admission rate decreased after 1982 (Figure 1). The decrease seems to have set in earlier and to have been more pronounced for patients undergoing an operation during their hospital stay. Figure 2 shows the admission rates of four age groups as a percentage of the admission rate in 1982, the year before budgeting was enacted (actual admission rates are available on request). For each group, the slopes of the regression lines for the periods before and after 1983 were determined (Table 1). They represent the average annual increase or decrease in the number of admissions in these periods. The differences between these slopes and their 95% confidence intervals were calculated. After 1983, the slopes of the lines for the admission rates decreased significantly in each of the four age groups. For patients 65 and over, however, the rate continued to rise. The group of patients between 45 and 64 displays an upward tendency up to 1983 but a decline afterward. The 15 through 44 age group had already exhibited a decreasing admission rate before 1983, a tendency that becomes more pronounced afterward. The admission rate of the youngest group, from 0 to 14 years, began to decline after 1983.

Length of Stay

In the past 20 years, hospital stays have gradually become shorter (Figure 3). The slope of the curve shows no change in 1983.

Patient Days and Daily Cost

Decreased admission rates, combined with shorter stays, have led to a spectacular drop in the number of patient days (Table 2): in 1988 the number of patient days was more than one fifth lower than would be expected by applying 1982 age-specific rates to the 1988 population. About 75% of the decrease was due to a shorter length of stay and 25% to lower admission rates.

TABLE 1—Slopes of the Regression Lines of the Annual Admission Rate for 1972–1982 and 1983–1988 and Their Differences, by Age

Period	Age 0–14	15–44	45–64	65+
1972–1982	−0.36	0.48	1.79	5.37
1983–1988	−2.14	−3.11	−2.34	2.97
Difference	1.78	3.59	4.13	2.40
95% confidence interval	0.97, 2.60	2.85, 4.51	3.76, 4.46	1.36, 3.46

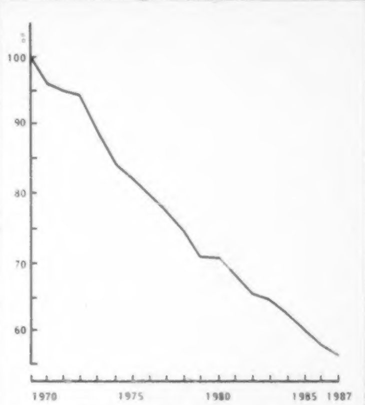

FIGURE 3—1969 age-adjusted average length of stay (18.0 days) divided by comparable length of stay for each year (1969–1988) and multiplied by 100.

Number and Rate of Operations

Table 3 shows that the number of operations performed on inpatients has decreased since 1977. The proportion of patients to undergo an operation has declined as well, and more steeply so since 1982. Figure 4 shows operation rates of the four age groups in percentages of the 1982 rate (actual rates are available on request). Table 4 presents the slopes of the regression lines. After 1983, the slopes of the regression lines were found to have decreased significantly in the age groups 0 through 14, 15 through 44, and 45 through 64, but not in the 65 and over group. For the 0 through 14 and 15 through 44 age groups, the decrease in the operation rates became more pronounced after 1982. For the group of 45 through 64-year-olds, the slight increase in operation rate before 1982 reversed to a slight decrease after that year. The operation rate of the oldest patient group, 65 and over, has, on the contrary, increased steadily.

Type and Weight of Operations

The demanding nature of operations has increased from a weight of 4.1 in 1978/1979 to 4.5 in 1985/1986. Table 5 lists the 20 operations most frequently performed on patients 65 and over in 1986, relating their frequency and age-adjusted rate in 1985/1986 to those in 1977 through 1979 for age groups below 65 and 65 and over. For all except three of these procedures, there has been an increase in frequency for the older age group; the incidence appears to have increased at a faster rate for the older age group than for the younger.

Hospital Mortality

Figure 5 expresses inpatient death rates in successive years in percentages of the 1977 death rates, separately for patients who did and did not undergo surgery while in the hospital and for age groups, younger and older than 65. The death rate appears to have gradually decreased for all four groups, with no observable effect on the budgeting requirement.

Discussion

The decline in admission rates and the decrease in the length of stay, which had already begun prior to 1982, indicate that hospitals had been trying to become more efficient for a number of years. In the past few years, an increasing proportion of operations have been performed in outpatient clinics or as day surgeries, the increase being as much as 31% between 1982 and 1988.[2] The shift is probably due to the spread of new diagnostic and therapeutic modalities applicable to outpatients and home treatment. However, in addition to improvements in operative techniques, many surgeries, such as extraction of the lens, are being done more frequently on less severely ill patients. Physicians as well as society at large have changed their attitude toward the necessity of hospital treatment for many medical conditions.

Although average length of stay had already decreased and inpatient operation rates for persons between 0 and 14 years and between 15 and 44 years, as well as the admission rate of 15 through 44 year olds, had declined even before budgeting was introduced, budgeting seems to have reinforced these tendencies. In the new budgeting system, outpatient treatment is more profitable for hospitals. However, with the fee-for-service system for medical specialists, inpatient treatment yields more income for them. Nevertheless, medical specialists know that if the expenses of a certain year exceed the budget, the budget for their hospital for the following year will be decreased; They have been willing to shift many procedures to the outpatient setting or to day surgery, a shift that involves only patients under the age of 65. We conclude that medical specialists have made well-considered decisions, probably motivated in part by guidelines and protocols. In the time period studied, there was no change in the fee of medical specialists for day surgery or outpatient care in relation to inpatient treatment. Although the hospital administration may have tried to influence a shift toward day surgery, it is the medical specialists who make the ultimate decision. The care of patients 65 and over has intensified, so the drive for efficiency has not been made at the expense of the oldest age group. Indeed, the fact that such additional care to aged patients could be provided within the budgets of the last few years is due to the more efficient treatment of younger patients.

The increasing complexity of inpatient operations, the rising number of operations performed on patients 65 and over, and the presumed poorer health of inpatients might be expected to increase mortality. However, this has not occurred. It is possible that, because of shorter hospital stays, some of the mortality has been shifted to the period after discharge. This does not seem to have occurred in the years involved, however; the proportion of inpatients discharged to a nursing home declined, the percentage of deaths occurring at home did not change, and the country's total death rate (8.4 per 1000 inhabitants in 1988[4]) declined.[5] Moreover, many of the operations most often performed on members of the older age group, such as lens extraction and total hip replacement (the very types of surgery that increased most in number), are aimed not so much at extending life expectancy as at enhancing the quality of life. Health care of the elderly has become more effective.

In spite of the growing population and an increase in the numbers of aged persons in the population, the government

TABLE 2—Total Number of Patient Days in 1982–1988, Number Expected on the Basis of the Rates of Patient Days by Age and Sex Groups in 1982, the Difference between the Two, and Cost per Patient Day, Expressed in 1988 Dutch Guilders

	No. of Patient Days			Cost per Patient Day[a] in 1988 Dutch Guilders
Year	Real	Expected	Difference	
1982	20 536 235	—	—	537.41
1983	20 068 864	20 716 653	647 789	547.78
1984	19 211 466	20 955 940	1 744 474	549.63
1985	18 746 418	21 218 140	2 471 722	575.11
1986	18 226 623	21 490 083	3 263 460	604.77
1987	17 433 865	21 772 010	4 338 145	634.78
1988	17 154 190	22 041 618	4 887 428	653.86

Source: Last column: reference 2 and personal communication by Mr. J. U. Brouwer, senior staff member National Hospital Institute.

Note. The last column gives the costs per patient day corrected, in Dutch guilders, for inflation rates. From 1982 to 1988 the total inflation rate was 8.9%. The costs of day surgery and outpatient departments are apportioned among the number of patient days, that is, hospital costs for day surgery and outpatient care are apportioned to the costs per patient day.

[a]Exchange rate July 1990: 1 Dutch guilder = 0.54 US dollar.

TABLE 3—Age-Adjusted Number of Inpatient Operations per 100 000 Population in 1977–1988, and as a Percentage of the Total Number of Admissions in 1969–1988.

Year	Operations per 100 000 Population	Admissions Operated on (%)
1969–1976		48.2
1977	5283	47.7
1978	5380	47.5
1979	5417	48.6
1980	5357	47.1
1981	5266	46.4
1982	5298	45.6
1983	5206	46.0
1984	4957	45.1
1985	4756	44.0
1986	4617	43.6
1987	4412	43.4
1988	4390	43.1

TABLE 4—Slopes of the Regression Lines of the Annual Operation Rate for 1977–1982 and 1983–1988 and Their Differences, by Age

	Age			
Period	0–14	15–44	45–64	65+
1977–1982	−0.75	−0.36	0.45	2.35
1983–1988	−3.03	−2.00	−1.26	2.28
Difference	2.28	1.64	1.71	0.07
95% confidence interval	1.43, 3.12	0.86, 2.43	1.42, 1.99	−0.55, 0.68

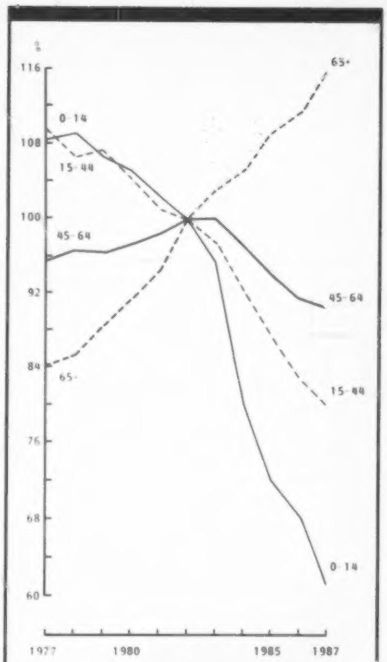

FIGURE 4—1982 inpatient operation rates for age groups 0–14 (44.1/1000 population), 15–44 (50.5/1000 population), 45–64 (55.0/1000 population), and 65 years and over (62/1000 population) divided by comparable rates for each year (1977–1988) and multiplied by 100.

was able to reduce the number of beds in general, short-stay hospitals from 69 612 to 65 510 in 1986,[6] a decrease from 4.5 to 4.2 beds per 1000 inhabitants. However, this last ratio is still higher than in most other developed countries with the exception of the Scandinavian countries.[7]

Obviously, there are more aspects of health care that have not been dealt with in this study, such as type and number of diagnostic procedures performed on inpatients and the way outpatient practices are conducted. Unfortunately, no reliable data are available on those aspects.

An accurate answer as to how much money has been saved, by the more efficient hospital use since 1983 is difficult to estimate. From Table 2 it can be calculated that the total hospital costs (number of patient days multiplied by costs per patient day) were about the same in 1982 and 1988: 11.0 and 11.2 billion Dutch guilders, respectively. However, this occurred in spite of about a 7% increase in expected patient days due to an increase in the aging population. So hospitals have been able to treat more older patients within the same budget. On the basis of demographic developments alone, the expenditures of hospitals should have been 7% more in 1988 than in 1982 (22 041 618 − 20 536 235 × 653.86), or nearly 1 billion guilders. But, the savings have been even greater because treatment of older patients is more expensive per patient day and because more expensive diagnostic and therapeutical procedures were increasingly used in hospitals during those years.

The drop in the hospital occupancy rate from 86% in 1982 to 73% in 1988 should have resulted in an even greater reduction in the number of hospital beds and, thus, in the long run, greater savings. But the total savings achieved so far are

TABLE 5—The 20 Operations Most Frequently Performed in 1986: Ratio of Age-Adjusted Rates in 1985–1986/1977–1979 for Persons Under 65 Years of Age and 65 and Over

Operation	Weight[a]	Ratio of Age-Adjusted Operation Rates for 1985–1986/1977–1979[b] 0–64	65+
Coronary bypass	7	3.77 (5375)	18.13 (2490)
Peripheral arterial bypass	7	1.62 (1434)	2.86 (2292)
Urethrotomy internal	1	1.16 (3306)	2.44 (1029)
Extraction of lens	5	1.59 (4287)	2.12 (14 656)
Transurethral resection of prostate	6	1.39 (3489)	1.66 (9429)
Total hip replacement	7	1.53 (2872)	1.64 (7600)
Resection of rectum	7	1.56 (1040)	1.48 (1677)
Segmental resection of colon	7	1.15 (1567)	1.38 (1814)
Insertion of cardiac pacemaker	6	1.26 (597)	1.38 (2425)
Transurethral resection of bladder	5	1.13 (3998)	1.33 (3145)
Femoral prothesis	7	1.05 (232)	1.32 (3266)
Operation of hip fracture	7	1.08 (1648)	1.31 (5652)
Abdominal hysterectomy	7	0.74 (11 735)	1.25 (1235)
Vaginal hysterectomy	7	0.73 (7553)	1.19 (1625)
Removal of osteosysthesis material	2	1.00 (10 566)	1.15 (1853)
Curretage of uterus	3	0.63 (17 912)	1.06 (2398)
Excision of lesions of skin	1	0.75 (6937)	1.03 (1498)
Exploratory laparotomy	5	0.83 (2319)	0.79 (1378)
Repair inquinal hernia	4	0.90 (16 535)	0.77 (5567)
Cholecystectomy	6	0.69 (8965)	0.75 (4989)

[a] 1 = easy, 7 = very demanding.

[b] The numbers of operations performed in 1986 are in parentheses.

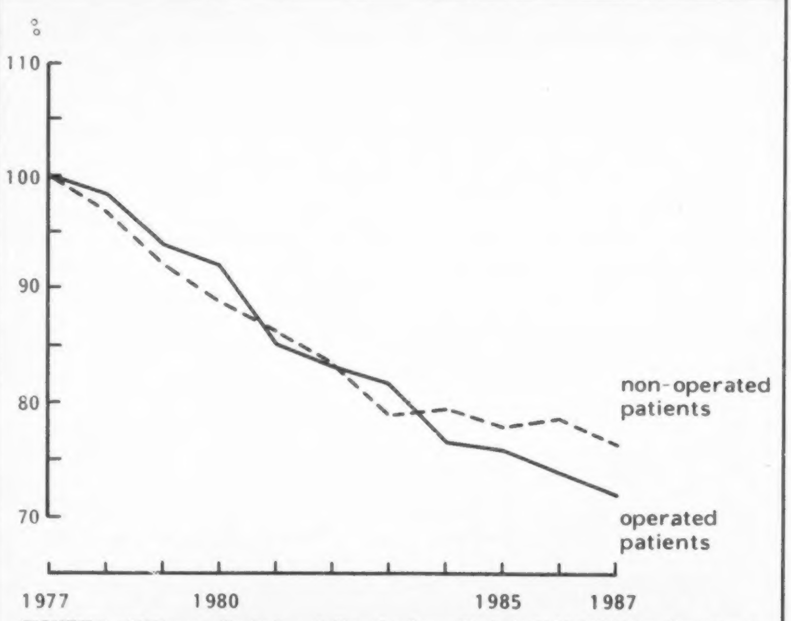

FIGURE 5—1977 age-adjusted mortality rates for patients 0–64 (4.53/1000 patients) and 65 years and over (60.6/1000 patients) undergoing an operation and patients 0–64 (21.2/1000 patients) and 65 years and over (163.5/1000 patients) not undergoing an operation divided by comparable rates for each year (1977–1988) and multiplied by 100.

already substantial. The total per capita costs of health care in the Netherlands, expressed in 1988 prices, which had increased continuously until 1983, have remained roughly at par since the enactment of budgeting.[6] The per capita costs were 2940 Dutch guilders in 1982 and 3000 Dutch guilders in 1988, or 10.3% and 9.9%, respectively, of the gross national product. The proportions of hospital costs were 59.7% and 57.8%, respectively, in these years. Indeed, the efficient use of inpatient hospital services has made it possible for other parts of the health care system, such as care for seniors, to be extended. But we must keep in mind that the population will continue to age, and that the cost of health care is bound to increase again unless the older age group, too, can be treated as outpatients or in day surgery and unless the treatment of the younger age groups can become even more efficient. Indeed, in the last 2 years, there have been indications that the financial constraints on hospitals have reached their limit or even have gone too far. In some hospitals, operating rooms and departments have been closed temporarily and, in May 1990, there was a strike by nurses because of the high work load and low income.

Conclusions

In recent years and, in particular, since the introduction of hospital budgeting in 1983, hospital services have been used more efficiently. That is probably the main reason why the rising tendency of total health care costs has leveled off since 1983. Additional care to aged patients could be provided within the budgets of the last few years because of the more efficient treatment of younger patients. Also, the decline in hospital mortality indicates that health care may have become more effective. Most operations carried out on senior patients are now aimed at improving the quality of life rather than lengthening it. Governmental measures to control the costs of health care have been accepted, so far, by health care providers. ☐

Acknowledgments

Parts of this study were published earlier in the *Nederlands Tijdschrift voor Geneeskunde*.

The authors wish to thank Prof. Chr. L. Rümke MD, PhD, retired professor of medical statistics, for his help with the statistical analysis. Mrs. R.R.M. de Groot, MD from the Centre for the Registration of Medical Data from Hospitals has made some final calculations.

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The Hazards of Extremely Low Frequency Radiation

The biological effects of extremely low frequency radiation (ELF) are unclear but "prudent avoidance" of exposure to electromagnetic radiation is advised, according to the American Industrial Hygiene Association (AIHA).

ELF, defined as radiation of 50 to 60 Hz from electric and magnetic fields, comes from appliances, wall wiring, ground currents, and service distribution lines—anything that produces electrical current. Electromagnetic radiation consists of electric radiation and magnetic radiation, but the magnetic radiation is considered to be the source of potential adverse effects.

The Environmental Protection Agency (EPA) has said that it can no longer categorically state that there is no effect from this radiation. Growing public concern about reported leukemia risk (especially in children), a possible connection with brain cancer, and potential interference with pacemakers has raised serious questions about the carcinogenicity of electromagnetic radiation.

According to ELF expert Otto White, CIH (Brookhave National Laboratory), former board coordinator to AIHA's Nonionizing Radiation Committee, several methods can mitigate ELF radiation effects. Appliance shielding, either with aluminum foil or more costly special alloys, can "attenuate magnetic fields" and limit exposures, White stated. Cancellation circuits, in which parallel systems of electric current cancel each other's effects, can also be used.

Romex cable and metal conduits can be used to prevent radiation from penetrating into the atmosphere from wiring. These methods "shield emissions so they don't go outside the device," White said. Grounding connections within a home or building can inadvertently create radiation emission zones. The proper design of grounding systems can eliminate this source, White noted.

Siting guidelines involving the locations of power lines, buildings, and residences are being increasingly examined to determine whether better locations can prevent exposures.

AIHA is a society of professionals dedicated to the health and safety of workers and the community. With more than 9000 members, AIHA is the largest international association serving occupational and environmental health professionals practicing industrial hygiene in private industry, academia, government, labor, and independent organizations.

Dogs as Sentinels for Lyme Disease in Massachusetts

Joann M. Lindenmayer, DVM, MPH, Durwood Marshall, BS, and Andrew B. Onderdonk, PhD

ABSTRACT

Background. An investigation of the relationship between incident human cases of Lyme disease and seroprevalence of antibodies to *B. burgdorferi* in dogs was undertaken in order to determine whether dogs might serve as sentinels for Lyme disease.

Methods. 3011 canine serum samples were analyzed by ELISA for antibody to *B. burgdorferi*. Records of incident human cases of Lyme disease were obtained from the Massachusetts Department of Public Health.

Results. Regression analyses of the relationship between the $\log_{10}$ (mean incidence in people 1985–1989) and canine seroprevalence from July 1988–August 1989 revealed that canine seroprevalence was highly predictive of incidence ($R^2 = 0.86$, $p < .0001$). A logistic regression model that incorporates the altitude of the town where each dog was resident, the date of sampling, and information on each dog's age, sex, and breed adequately explained the risk of canine seropositivity. Dogs resident at altitudes less than 200 feet, of sporting or large mixed breeds, and greater than two years of age were five times, four times, and almost three times more likely, respectively, to exhibit seropositivity than were other dogs.

Conclusions. Estimates of the prevalence of antibody to *B. burgdorferi* in dog populations offers a sensitive, reliable, and convenient measure of the potential risk to people of *B. burgdorferi* in the environment. Risk factors for canine seropositivity may directly or indirectly illuminate certain aspects of the epidemiology of human Lyme disease. (*Am J Public Health.* 1991;81:1448–1455)

Introduction

Summary statistics of incident human cases collected by the Centers for Disease Control and state departments of public health have been used to highlight national and regional distributions and changes in Lyme disease epidemiology.[1-3] Passive surveillance systems, however, are well known for underestimating actual cases of any disease. The implementation of measures designed to minimize human exposure to ticks and a geographically variable index of suspicion for Lyme disease may affect numbers of case reports, but they have little to do with changes in the environmental burden of *B. burgdorferi.*

Testing dog populations for the presence of antibody to *B. burgdorferi* has been used as an epidemiologic tool to supplement surveillance based on passive case reporting.[4-10] Dogs in endemic areas have been shown to be almost six times as likely as people to be exposed to infected ticks,[11] making them a sensitive indicator of the presence of *B. burgdorferi.* The present study was undertaken to determine whether a relationship exists between seroprevalence of antibodies to *B. burgdorferi* in the dog population and incident cases of human Lyme disease in Massachusetts. If so, dogs might serve as sentinel animals for Lyme disease.

Methods

Canine sera

Thirty-two veterinary practices located in eight Massachusetts counties participated in this study (Figure 1). Practices were chosen on the basis of the following criteria: routine use of Tufts Veterinary Diagnostic Laboratory (TVDL), willingness to participate in the study for the period of 1 year, and location. Practices were chosen to represent counties that were known, based on Massachusetts Department of Public Health records, to be endemic for Lyme disease (Dukes, Nantucket, Essex, Barnstable, and Plymouth counties), to represent areas where Lyme disease activity has recently been reported (Hampshire and Hampden counties), or to be located in areas where Lyme disease had not previously been reported (Worcester).

At each veterinary practice, veterinarians were asked to choose a convenient day of the week, and on that day to take serum samples from the first 5 to 10 consecutive dogs that were seen, regardless of the reason for presentation. If a client did not permit blood to be drawn, the veterinarian skipped that dog and continued to sample consecutively until the predetermined number of samples had been collected. For each dog sampled, veterinarians filled out brief forms providing information on the dog's breed, age, sex, owner's name and address, reason for presentation to the veterinarian, date the sample was obtained, and whether or not the dog was a Lyme disease suspect.

Sampling was undertaken from August 1, 1988, through December 31, 1988,

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Joann M. Lindenmayer is with the department of Comparative Medicine, Tufts University School of Veterinary Medicine. Andrew B. Onderdonk is Director, Microbiology Laboratory, Brigham and Women's Hospital, Boston, MA. Durwood Marshall is with Tufts University Academic Computer Services.

Requests for reprints should be sent to Joann M. Lindenmayer, DVM, MPH, Tufts University School of Veterinary Medicine, Department of Comparative Medicine, 200 Westboro Road, North Grafton, MA 01536.

This paper was submitted to the journal May 30, 1990, and accepted with revisions March 6, 1991.

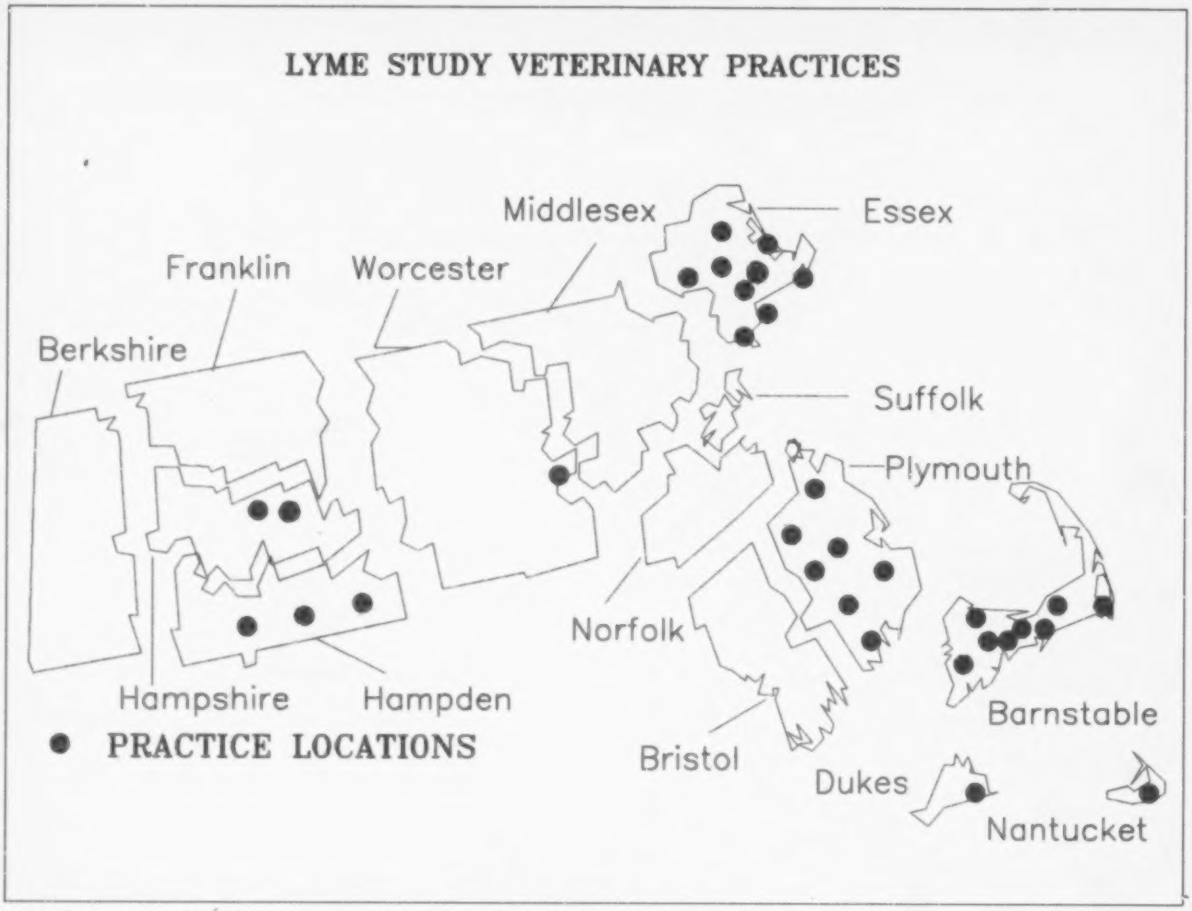

FIGURE 1—Lyme Study Veterinary Practices in Massachusetts Counties.

suspended for the winter, begun again on April 1, 1989, and continued through July 31, 1989. Samples were collected in serum separator tubes and sent to the TVDL where they were centrifuged and frozen at minus 20°C until testing could be performed. Serum samples from dogs suspected of having Lyme disease were tested within 7 days of arrival, and serum samples from dogs not suspected of having Lyme disease were tested within 4 months of arrival at TVDL. The presence of antibodies to *B. burgdorferi* was measured using an enzyme-linked immunosorbent assay (ELISA) as previously described.[12] The cutoff value for seropositivity of 0.170 optical density units is 3 standard deviations (SDs) above the mean optical density for serum samples taken from 34 apparently healthy dogs resident in areas not endemic for Lyme disease.

Human Lyme Disease Cases

Statistics on Lyme disease cases for the years 1985 to 1989 were obtained from the Surveillance Service of the Division of Epidemiology in the Massachusetts Department of Public Health.

Statistical Analysis

Canine seroprevalence was calculated as the number of dogs seropositive divided by the number of dogs tested for each category of interest. A logistic regression model was developed to estimate relative risks for canine exposure associated with five factors that were suspected, a priori, of being predictive of seropositivity.

1. *Altitude of the town where each dog was resident.* Several environmental factors have been proposed to explain the range of *Ixodes dammini.*[13–16] The distribution of incident cases of human Lyme disease in Massachusetts suggests an association between altitude, either directly or as a proxy for other factors, and incidence. Altitude was modeled on two levels depending on whether a town was located at <200 feet or ≥200 feet above sea level.

2. *Date the sample was drawn.* Nymphal ticks, the developmental stage most responsible for transmission of *B. burgdorferi* to people, have been demonstrated to be most active in late spring and early summer, whereas adult ticks have been most active in the fall and winter.[17] The sampling date was thus modeled on two levels, depending on whether a sample was obtained between August 1 and December 31, 1988, or between April 1 and July 31, 1989. These dates roughly correspond to fall/winter and spring/summer seasons.

3. *Sex.* This factor was modeled on two levels, male and female, regardless of whether a dog was intact or neutered/spayed. Previous studies indicate that male dogs may be predisposed to exposure to *B. burgdorferi.*[9]

4. *Breed.* Three levels of breed were chosen to reflect the relative probability of

TABLE 1—Comparison of Annual Incidence[a] of Human Lyme Disease with Canine Seroprevalence,[b] 1988–1989, by Massachusetts County

	Human						Dog	
	Year						Year	
County	85	86	87	88	89	$\bar{X}$[c]	88–89	(95% CI[d])
Barn	2.4	8.9	1.8	2.4	8.3	4.8	0.283	(0.252–0.314)
Berk	—	0.7[e]	—	0.7[e]	2.8	0.8	—	
Brist	—	0.6	—	2.9	5.0	1.7	0.044	(−0.016–0.104)
Dukes	30.1	40.2	391.8	—	20.1	96.4	0.495	(0.398–0.592)
Essex	0.8	3.2	1.6	1.4	2.5	1.9	0.174	(0.147–0.201)
Frank	1.5[e]	—	—	1.5[e]	3.0	1.2	0.077	(−0.025–0.179)
Hampden	0.2[e]	1.4	0.5	1.6	1.6	1.1	0.041	(0.013–0.069)
Hshire	0.7[e]	2.8	0.7[e]	1.4	6.9	2.5	0.051	(0.007–0.095)
Mid	0.1[e]	1.3	1.1	1.3	1.7	1.1	0.016	(−0.006–0.038)
Nant	—	378.0	31.5	—	—	81.9	0.691	(0.614–0.768)
Norfolk	—	1.3	1.2	1.6	1.0	1.0	0.029	(−0.011–0.069)
Ply	0.7	1.2	—	0.7	2.8	1.1	0.138	(0.109–0.167)
Suffolk	0.2[e]	1.3	1.6	0.3	0.2[e]	0.7	0.125	(−0.104–0.354)
Worc	0.2[e]	0.3	0.9	1.1	1.2	0.7	0.013	(−0.012–0.038)

Note. Barn = Barnstable; Berk = Berkshire; Brist = Bristol; Frank = Franklin; Hshire = Hampshire; Mid = Middlesex; Nant = Nantucket; Ply = Plymouth; Worc = Worcester;— = no cases reported/no dogs tested.

[a] Cases/100 000 persons (Source: Massachusetts Department of Public Health).

[b] Number seropositive/number tested × 100.

[c] Mean incidence, 1985–1989.

[d] CI = confidence interval.

[e] Represents a single case.

exposure to ticks in the environment—sporting and large mixed breed dogs (high probability of exposure), toy and small mixed breed dogs (low probability of exposure), and all others, including working, hound, terrier, nonsporting, medium mixed breeds, and dogs of unknown breed status (intermediate probability of exposure).

5. *Age of the dog at the time of sampling.* The risk for dogs of testing positive for antibody to *B. burgdorferi* has been found to be associated with the cumulative time of *I. dammini* exposure,[11] although other studies have reported a greater degree of seropositivity and higher antibody titers among younger than among older dogs.[6,9] Age of dogs, as a proxy for cumulative time of exposure, was modeled on two levels, ≤2 years of age and >2 years of age.

Logistic regression analyses were carried out with BMDP and SAS, and Pearson χ^2 analysis was performed with StatXact (Cytel Software Corporation, Cambridge, Mass).

Results

Characteristics of the Study Population

During the period August 1, 1988, to July 31, 1989, 3011 serum samples obtained from 2941 Massachusetts dogs were tested for the presence of antibody to *B. burgdorferi.* Seventy dogs, or 2% of the total population, were sampled twice, the second time for reasons unrelated to the present study. Males and females were approximately equally represented (46.4% vs 49.5%, respectively). Sex of the remaining 4.1% of the study population was not noted by veterinarians. Sporting, large mixed breed, and working dogs comprised the majority of the breeds represented (29.4%, 24.3%, and 19.6%, respectively). All other breeds were approximately equally well represented. The percentage of dogs in each age category ranged between 5.7% and 8.9% up to 11 years of age. Above that, the percentage in each age category ranged between 0.1% (>15 years old) and 4.7%.

Seroprevalence

Among all dogs tested, unconditional seroprevalence was 20.3%.

Comparison of incidence and prevalence. Canine seroprevalence was calculated by dogs' place of residence, the most likely site of exposure to *B. burgdorferi.* This permitted inclusion of 262 of 351 Massachusetts communities in the analysis.

Great variation exists among incident human cases by county from year to year (Table 1). By contrast, canine seroprevalence data were available for every county in the state except one, and, because so many dogs were sampled, reasonable confidence intervals could be calculated for each point prevalence estimate. A comparison of incidence and prevalence by town for a single county (Essex) likewise confirms the completeness and stability of the canine data (Table 2).

Regression analyses of the relationship between canine seroprevalence in 1988 to 1989 and human incidence per 100 000 persons for each year between 1985 and 1989 by county were unable to statistically confirm the utility of canine seroprevalence as a predictor of the number of human cases. This is likely attributable to the small number of documented county-specific cases in any given year. In order to overcome the variability inherent in the human data, cases of Lyme disease were averaged by county for the years 1985 through 1989. When the relationship between $\log_{10}$ (mean incidence in humans 1985–1989) and canine seroprevalence was explored, canine seroprevalence was found to be highly predictive of incidence ($R^2 = 0.80$, $P < .0001$) (Figure 2).

Seroprevalence by month. Seroprevalence by month of sample submission did not vary with the exception of November 1988, when a particularly high proportion of samples exhibited seropositivity (Figure 3).

TABLE 2—Comparison of Annual Incidence of Human Lyme disease[a] with Canine Seroprevalence,[b] 1988–1989, in towns of Essex County, Massachusetts

	Human					Dog	
	Year					Year	
Town	85	86	87	88	89	88–89	(95% CI[c])
Andover	3.7[d]	—	—	—	3.7[d]	0.029	(−0.027–0.085)
Beverly	—	2.6[d]	—	5.3	2.6[d]	0.057	(−0.020–0.134)
Boxford	—	—	17.8[d]	—	—	0[e]	
Danvers	—	—	—	—	—	0.088	(−0.007–0.183)
Essex	—	—	—	—	65.8	0.677	(0.563–0.791)
Georgetown	—	—	—	—	—	0.167	(−0.044–0.378)
Gloucester	—	—	—	—	—	0.294	(0.186–0.402)
Groveland	—	—	—	—	18.9[d]	—	
Hamilton	—	—	—	—	14.5[d]	0.452	(0.277–0.627)
Haverhill	—	—	—	—	—	0[e]	
Ipswich	35.4	123.9	44.2	26.5	35.4	0.350	(0.202–0.498)
Lawrence	—	1.6[d]	—	—	—	—	
Lynn	—	—	1.3[d]	—	—	0.077	(−0.025–0.179)
Lynnfield	—	—	—	—	—	0.083	(−0.073–0.239)
Magnolia	—	—	—	—	—	0.333	(0.025–0.641)
Manchester	—	—	18.6[d]	—	—	0.100	(−0.031–0.231)
Marblehead	—	—	—	—	—	0[e]	
Methuen	—	—	—	—	—	0[e]	
Middleton	—	—	—	—	—	0.043	(−0.040–0.126)
N. Andover	—	—	—	—	—	0.091	(−0.079–0.261)
Nahant	—	—	—	—	—	0[e]	
Newbury	—	21.1[d]	—	—	—	—	
Newburyport	—	—	—	12.1	—	0.200	(−0.151–0.551)
Peabody	—	—	—	—	4.5	0.024	(−0.023–0.071)
Rockport	—	—	—	—	31.5	0.263	(0.065–0.461)
Rowley	—	—	—	—	—	0.333	(−0.044–0.710)
S. Hamilton	—	—	—	—	—	0.222	(0.065–0.379)
Salem	—	2.7[d]	2.7[d]	—	—	0.031	(−0.029–0.091)
Salisbury	—	—	—	—	—	0.200	(−0.151–0.551)
Saugus	—	—	—	—	—	0[e]	
Swampscott	—	7.3[d]	—	14.6	—	0[e]	
Topsfield	—	—	—	—	—	0.167	(−0.005–0.339)
W. Peabody	—	—	—	—	—	0[e]	
Wenham	—	25.8[d]	25.8[d]	—	51.7	0.200	(−0.151–0.551)

Note. — = no cases reported/no dogs tested.

[a] Cases per 100 000 persons (Source: Massachusetts Department of Public Health).

[b] Number seropositive/number tested × 100.

[c] CI = confidence interval.

[d] Represents a single case.

[e] Represents testing of ≥5 dogs.

Seroprevalence by breed. Seropositivity was highest among sporting and large mixed breed dogs and lowest among toy and small mixed breed dogs (Figure 4).

Seroprevalence by age. Degree of seropositivity was lowest among dogs ≤2 years of age. Above 2 years of age, seropositivity remained relatively stable (Figure 5).

Sensitivity of the Canine Model

Using seropositivity alone, far more dogs tested positive for exposure to *B. burgdorferi* than incident human cases were reported. Comparing mean incidence for 1985 to 1989 with canine seroprevalence by county, seropositivity estimates were 513 to 17 857 times greater than incident human cases.

Veterinary suspicion for Lyme disease was associated with serologic test results. Of 203 dogs that were noted to be Lyme disease suspects based on clinical presentation, 91 (44.8%) tested positive for exposure to *B. burgdorferi*, whereas only 18.4% (465/2530) of nonsuspects and 20.1% (56/278) of dogs whose Lyme disease status was not noted tested positive for exposure. The Pearson χ^2 test statistic was 141.7 (df = 2, P = .0011) and was based on a Monte Carlo simulation of 2000 tables generated from fixed marginal distributions of the observed data.

Logistic regression. Of the dogs studied, 2789 met the criteria that comprise the population profiles for the logistic regression model and were included in the analysis. The standard logistic regression model had the following form:

$$\ln(p/1-p) = -4.74 + 1.64T - 0.20D - 0.05S + 0.82B1 + 1.41B2 + 1.03A,$$

where

p = probability of testing positive
1−p = probability of testing negative
T = altitude of town where dog was resident
D = season serum sample was obtained
S = sex of dog
B1 = indicator variable for the differential effect of moderate vs low exposure by breed
B2 = indicator variable for the differential effect of high vs low exposure by breed

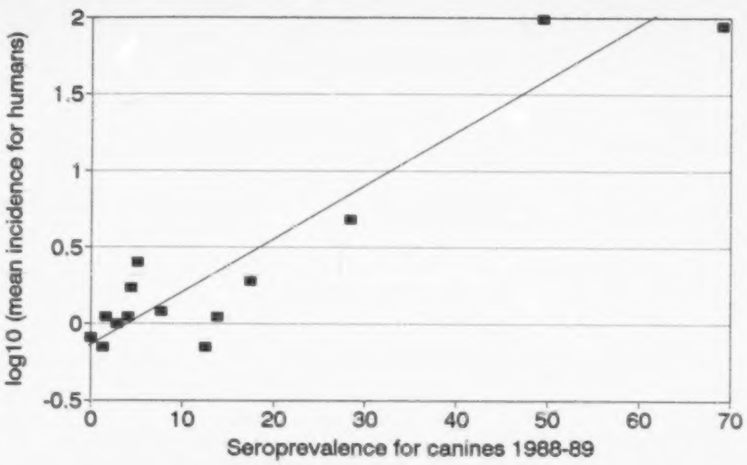

FIGURE 2—Least-Squares Regression Line for $\log_{10}$ Mean Intercept = 0.13, Slope = 0.03, Adjusted $R^2 = 0.86$ ($P < .001$). Human Incidence 1985–89 vs Canine Seroprevalence, for 13/14 Massachusetts Counties.

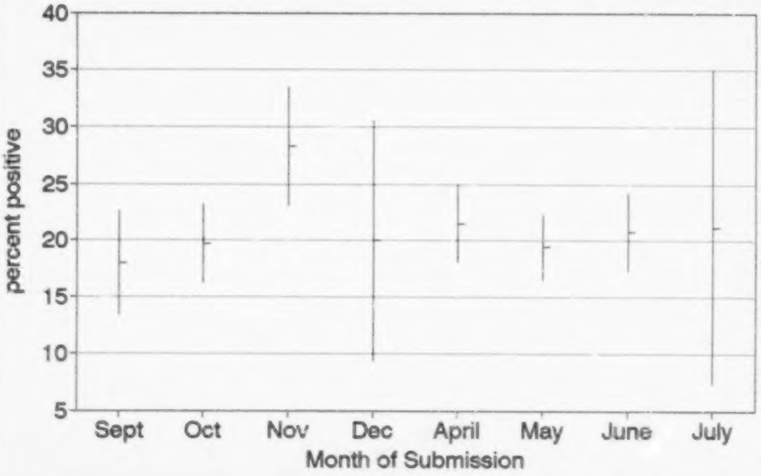

FIGURE 3—Canine Seroprevalence by Month (mean ± 95% confidence interval).

TABLE 3—Odds Ratios (OR) for Logistic Regression Factors

Factor	OR(95% CI[a])
Intercept	0.01 (0.00–0.02)
Altitude	5.15 (3.37–7.87)
Date	0.82 (0.67–1.00)
Sex	0.95 (0.79–1.15)
Breed1[b]	2.28 (1.23–4.23)
Breed2[c]	4.10 (2.23–7.51)
Age	2.80 (2.01–3.91)

[a] CI = confidence interval

[b] Breeds with moderate probability of exposure.

[c] Breeds with high probability of exposure.

A = age of dog at time of sampling

The χ^2 goodness-of-fit statistic for the equation sum[2 $O_i \times \ln(O_i/E_i)$] was 34.37 (df = 39, $P = .68$), suggesting that the model adequately explained the test outcome. The 48 possible cells resulting from the coding of the five factors yielded two cells without any observed data due to incomplete records. Two of the 46 remaining cells proved to be particularly unusual and not well accommodated by the logistic model. These two cells were low-exposure breed female dogs <2 years of age, resident at low altitudes and sampled in the spring or summer (n = 3), and low-exposure breed male dogs ≥2 years of age, resident at high altitudes and sampled in the fall or winter (n = 5). These cells, because of their small counts, are unlikely to be representative of the corresponding population profiles. For this reason, they were removed from the analysis. The recalculated χ^2 goodness-of-fit statistic was 29.84 (df = 37, $P = .825$). All final results are based on the latter model.

Odds ratios for factors included in the model (Table 3) revealed that dogs resident at or near sea level were approximately five times as likely to test positive for exposure as dogs resident at higher elevations. Breeds with moderate probability of exposure and those with high probability of exposure were more than two times and more than four times as likely, respectively, to test positive than breeds with low probability of exposure. Linear contrast analysis for differences between breeds with high probability of exposure and all others indicated that the former were far more likely to be exposed than the latter ($\chi^2 = 33.01$, $P = .0001$). Dogs >2 years of age were almost three times as likely to test positive as dogs ≤2 years of age. Dogs sampled in the spring or summer were slightly less likely than dogs sampled in the fall or winter to test positive, although this difference was not significant. Sex of the dog had no effect on the probability of testing positive.

In the logistic regression model, cells with the highest probabilities for seropositivity (>0.22) represented dogs that lived at altitudes <200 feet, of the sporting or large mixed breed types, and older than 2 years of age. No differences with respect to date of sampling or sex were noted. Cells with the lowest probabilities for seropositivity (<0.017) represented dogs that lived at altitudes ≥200 feet, were <2 years of age, and were of low and moderately exposed breed types. As before, no differences in date of sampling or in sex were noted.

Discussion

The dog has been proposed for use as a sentinel animal to detect the presence of *B. burgdorferi* for two reasons. Dogs exposed to infected ticks develop antibodies to the spirochete and often exhibit clinical signs of the disease similar to those in people. Second, dogs are more likely to be exposed to infected ticks because they come into more physical contact with tick habitats than do most people, and because

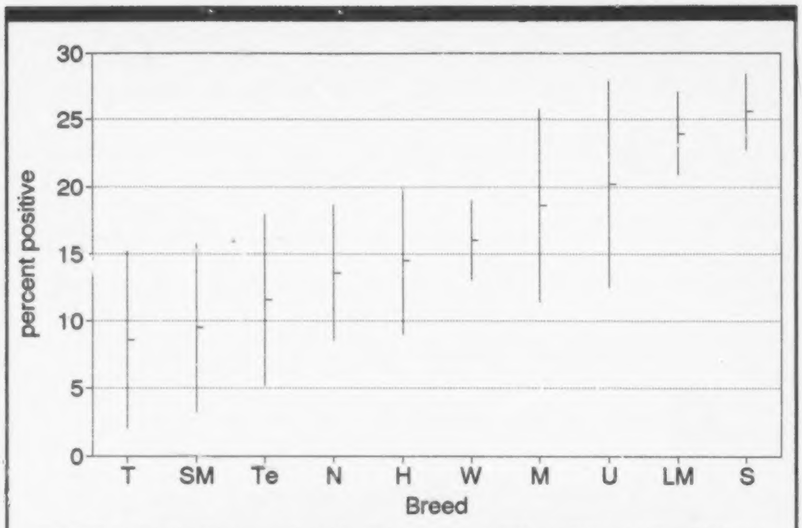

FIGURE 4—Canine Seroprevalence by Breed (mean ± 95% confidence interval; T=toy, SM=small mixed, Te=terrier, N=nonsporting, H=hound, W=working, M=mixed, U=unknown, LM=large mixed, S=sporting).

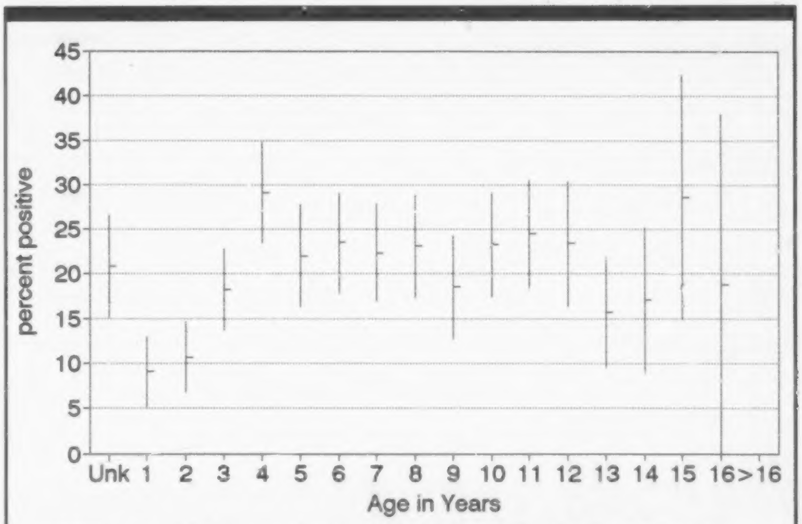

FIGURE 5—Canine Seroprevalence by Age (mean ± 95% confidence interval).

measures to prevent exposure to ticks are difficult to implement in dogs. Results of the present study indicate that under certain conditions canine seroprevalence estimates are highly predictive of incident human cases, further strengthening the usefulness of dogs as sentinel animals.

The present comparison between two different measures of disease and two different time periods is unavoidable for two reasons. Incident canine cases of Lyme disease are difficult to determine. Efforts to induce seroconversion to *B. burgdorferi* or clinical Lyme disease in laboratory dogs have not been uniformly successful.[18,19] Furthermore, dogs often exhibit seropositivity to the spirochete without concomitant clinical evidence of disease.[6–8,11,18–21] It is unlikely, then, that incident cases among dogs will ever be available for comparison with human incidence rates. Second, regression analyses of canine seroprevalence in 1988–1989 and human incidence for each year between 1985 and 1989 failed to reveal a relationship between the two. This is probably due to the great variability inherent in incident human cases from year to year. Seroprevalence estimates among dogs for a given year are likely to be stable because so many dogs can be sampled under an active surveillance system. It would, however, be useful to extend these studies over a longer period of time in order to determine whether or not these observations are maintained.

The overall seroprevalence estimate of 20.3% for this study does not reflect seroprevalence among all Massachusetts dogs because participating veterinary clinics were chosen to afford maximal coverage of particular counties in Massachusetts. For this reason, county- and town-specific seroprevalence rates are more likely than state-wide estimates to be representative of dogs presented for veterinary care in these regions. However, all seroprevalence estimates may actually underestimate true seroprevalence because they were based on dogs presented to veterinary practices rather than the general canine population. Inclusion of stray dogs and dogs that do not receive regular veterinary care might well increase seroprevalence estimates because these dogs would be more likely to run free and therefore be exposed to infected ticks.

Seroprevalence estimates were based on the assumption that seropositive dogs received exposure in the towns where they resided. In fact, dogs often travel with their owners, particularly in the summer months. This raises the possibility that individual seropositive dogs were exposed to infected ticks at locations other than their towns of residence and might account for the presence of seropositive dogs in towns where Lyme disease is not known to be endemic.

Previous studies in this laboratory and elsewhere[12,22–24] have documented false positive ELISA test results for both canine and human serum samples. While we recognize that this possibility exists, test results from an area where Lyme disease is not known to be endemic by Massachusetts Department of Public Health standards (i.e., Worcester County) indicate that a very small proportion of dogs (1/78 or 1.3%) tested positive for antibody to *B. burgdorferi*, presumptive evidence that the proportion of false positive samples in this study is low.

Use of the canine to detect and quantify Lyme disease risk offers several advantages over the use of human incidence rates alone. Dog populations are particularly convenient to test for exposure to *B. burgdorferi*. Dogs are routinely brought to veterinarians in the spring for heartworm testing. Blood drawn for this purpose can likewise be made available for Lyme disease testing, obviating the need for a separate visit to the veterinarian. Furthermore, dogs tested at this time of year are usually well animals presented for routine checkups and are thus more likely than dogs presented at other times of the year to be representative of veterinary clinic catchment populations.

Second, seroprevalence estimates in dogs not only reflect incidence rates in people, but they also appear to be a far more sensitive indicator of the presence of *B. burgdorferi* than do the latter. Seropositivity for every 100 000 dogs tested was consistently much higher than Lyme disease incidence per 100 000 persons.

Third, seroprevalence estimates in dogs are more likely to reflect the actual environmental risk of Lyme disease than are reports of incident human cases. Factors such as increasing awareness of Lyme disease may falsely increase incidence rates, whereas the use of measures to prevent exposure to ticks may actually decrease incidence rates apart from any changes in the environmental burden of *B. burgdorferi.* Measures designed to prevent dogs from exposure to potentially infected ticks are frequently unrewarding. Dogs' propensity to pick up large numbers of ticks and the difficulties associated with finding tiny *I. dammini* make tick exposure a near-certainty in endemic areas except, perhaps, for dogs that rarely venture outdoors.

Fourth, canine seroprevalence estimates permit a fairly complete representation of areas of potential risk to be drawn. As evidenced by the Massachusetts Department of Public Health data, collected under a passive reporting system, incidence measures for a particular county or town are frequently based on a single case report. Furthermore, marked year-to-year variability in incidence rates is frequently observed. Under an active canine surveillance system such as the one used here, additional information was made available to permit the calculation of point prevalence estimates and reasonable confidence intervals for 31 of 34 towns in one particular county (Essex) and for 13 of 14 counties.

The logistic model employed in the present study appears to adequately explain the risk of seropositivity among Massachusetts dogs and has defined certain risk factors for canine exposure, namely, altitude of town of residence, breed of dog, and age. Dogs that live at altitudes <200 feet above sea level appear to be at greater risk for exposure to *B. burgdorferi.* The fact that seroprevalence varies widely even at this altitude raises the possibility that ecologic factors such as humidity or temperature may affect deer tick survival and thus the risk of exposure to Lyme disease. Lyme disease is also known to occur in noncoastal areas of the United States where altitude may not be definable as in the present model.

Previous studies have not been able to confirm a breed disposition among seropositive dogs, although mixed breed dogs and hunting breeds were more frequently seropositive.[9] The present study, perhaps because of a larger sample size, establishes that seropositive dogs are more likely to be sporting and large mixed breed dogs. These animals are presumably at greater risk of exposure because the former are used for hunting and sporting purposes and the latter, by virtue of their size, may be allowed more freedom to roam.

Dogs >2 years of age appear to be at greater risk of exposure than younger dogs. Other studies have documented maintenance of antibody titers in dogs months to years after initial exposure to the spirochete.[20,21] Under these circumstances, seroprevalence could be expected to increase with advancing age. In the present study, however, age-specific seroprevalence estimates appear relatively stable above 2 years of age. It may be that, for a given dog population, only a fixed proportion of its members are at risk for exposure to *B. burgdorferi.* For example, a smaller proportion of dogs might be at risk in areas with large urban populations and strict leash laws than in rural areas with less restrictive leash laws. Dogs at risk may be exposed to infected *I. dammini* and seroconvert at an early age, and thus, after the critical age of 2 years, no increase in age-specific seropositivity would be noted.

In the present study, sex appeared not to explain seropositivity. Information on reproductive status of an animal, whether intact or neutered/spayed, was requested but frequently not provided. It is possible that reproductive status, as an indicator for proclivity to roam, might well be a risk factor for seropositivity, although this could not be determined in the present study.

The risk factors identified for dogs may directly or indirectly illuminate certain aspects of the epidemiology of human Lyme disease. Altitude can certainly be a risk factor for people. Age, as a proxy for cumulative duration of exposure, and breed, as a proxy for occupational or recreational exposure, are other factors that should be considered in models investigating the risk of Lyme disease in people.

A vaccine against *B. burgdorferi* has recently been made available for use in dogs (Borrelia burgdorferi Bacterin, Fort Dodge Laboratories, Fort Dodge, Iowa). Widespread use of such a vaccine could well compromise the ability of canine populations to serve as sentinel animals for Lyme disease. Vaccinated dogs would presumably be protected against field exposure to the spirochete and antibody titers in vaccinated dogs would be indistinguishable from those of naturally exposed dogs. The degree to which the dog-owning public avails itself or the vaccine remains to be seen. However, the vaccine is likely to be put to greatest use in Lyme disease–endemic areas where a sentinel system would be of least benefit. Dogs in nonendemic areas might well still serve as an early warning system for detection of the spirochete if they remain largely unvaccinated. ☐

Acknowledgments

This work was supported by a public health fellowship from the Massachusetts Health Research Institute of the Medical Foundation, Boston, Mass, to Dr. Lindenmayer.

The authors are grateful to the veterinarians who participated in this study, and in particular to Dr. Frederick Johnson, Dennis Animal Hospital, Dennis, and Drs. James Dasbach and John Rae, Nantucket Animal Hospital, Nantucket, Massachusetts, for their many helpful comments and suggestions. Mr. Paul Etkind, Director, Epidemiology Program of the Massachusetts Department of Public Health, critically reviewed the manuscript.

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APHA's Food and Nutrition Section Announces Awards

The Food and Nutrition Section of the American Public Health Association announces the availability of the Helen R. Stacey and Joseph A. Walsh awards for graduate education at the master's degree level in public health nutrition. Three awards are given consisting of $2000 each. These awards must be used for the academic year 1992 to 1993. Applications must be submitted no later than May 1, 1992.

Nominations are being solicited for the Mary C. Egan Award for Outstanding Professional Contributions and Service in Public Health Nutrition. This award carries with it a $1000 honorarium. Nominations must be submitted no later than May 1, 1992.

Applications and nominations forms may be secured from Elvira Jarka, MS, RD, Chair, Awards Committee, F&NS, APHA, Lake County Health Department, 2303 Dodge, Waukegan, IL 60085.

National Health Objectives for the Year 2000: The Demographic Impact of Health Promotion and Disease Prevention

Michael A. Stoto, PhD, and Jane S. Durch, MA

ABSTRACT

Background. The national objectives in *Healthy People 2000,* drafted by health professionals aware of currently available public health interventions, represent a wealth of information about near-term future mortality and morbidity.

Methods. Life table methods were used to calculate the impact of projected changes in mortality and activity limitation rates on life expectancy and expected disability years.

Results. Meeting the mortality objectives would increase life expectancy at birth by 1.5 to 2.1 years, raising life expectancy to 76.6 to 77.2 years. In addition, meeting the target for disability from chronic conditions would increase the number of years of life without activity limitations from 66.8 years to 69.3–69.7 years. If the targets for coronary heart disease and unintentional injury were changed to reflect recent trends, a greater improvement in life expectancy at birth would be achieved: from 1.8 to 2.7 years to 76.9 to 77.8 years.

Conclusion. Meeting the targets would have an important demographic impact. Including changes in the coronary heart disease and injuries targets, life expectancy in the year 2000 would be above the middle of the ranges used in current Census Bureau projections. (*Am J Public Health.* 1991;81:1456–1465)

Introduction

On September 6, 1990, the US Public Health Service (PHS) released *Healthy People 2000: National Health Promotion and Disease Prevention Objectives,*[1] a report on some 300 national health objectives for the year 2000. Based on the input of scientists and health professionals in many fields, PHS set specific numerical targets that are thought to be achievable over the next decade and identified data to track progress toward these targets. As experience with the earlier 1990 Objectives for the Nation[2] has shown, tracking these objectives can help to identify areas of success as well as areas that need more attention and resources. Thus the objectives will help guide health policy over the next decade.

The objectives are organized into 22 "priority areas" relating to specific problems, conditions, or diseases; risk factors; or potential interventions (see Table 1). Each priority area presents objectives for particular health status measures, for reduction of risk factors, and for provision of health services and protective measures. In total, the report contains about 300 objectives for the general population plus more for minority groups and other populations of special concern.

The objectives bring to bear in a specific and quantitative way the informed judgment of public health experts on a critical demographic issue, future improvements in mortality and morbidity. The objectives were developed by groups of scientists and public health professionals who are aware of the currently available and soon-to-be available public health interventions. Therefore, they represent a wealth of information about both future mortality and morbidity in the United States and the potential for health promotion and disease prevention interventions.

The objectives were, however, written by many independent committees and appear in a number of different formats. Some objectives target improvements in specific causes of mortality and morbidity but use noncomparable or nonstandard measures to provide a baseline. Other objectives specify potential changes in risk factors, use of preventive services, and other implementation strategies, and they are not directly translated into mortality or morbidity measures.

The intent of this paper is to organize and analyze in demographic terms the potential for future reductions in mortality and morbidity implicit in the national objectives. Specifically, we calculate the impact on life expectancy and other summary demographic measures of meeting the targets set in the objectives. This analysis serves two purposes. First, it summarizes in quantitative terms the scientific and medical knowledge of the many experts who participated in setting the objectives. This should be of interest to demographers and forecasters who want a stronger scientific backing for their future mortality and morbidity projections. Second, it offers guidance for enhancing further efforts in health promotion and disease prevention. The analysis shows where the potential for health status improvements is high and low and thus in-

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The authors are with the National Academy of Sciences in Washington, DC.

Requests for reprints should be sent to Michael A. Stoto, PhD, Institute of Medicine, National Academy of Sciences, 2101 Constitution Avenue, Washington, DC 20418.

This paper was submitted to the journal December 26, 1990, and accepted with revisions February 19, 1991.

forms discussions about setting priorities among the objectives.

Methods

A number of technical features of objectives development should be taken into account in trying to develop measures of the overall mortality impact of meeting the objectives.

Because the objectives were drafted by 21 different working groups, the statistical measures used lack the unity of method that they would have if they had been developed by demographers and statisticians. Among mortality measures, for instance, nonstandard cause-of-death categories, rather than the categories used in vital statistics reports[3] and *Health, United States 1989,*[4] are used for coronary heart disease, diabetes, pneumonia and influenza, and a number of specific kinds of injuries.

One of the hallmarks of the national objectives process is that it sets specific numerical targets for the future. This allows policymakers to assess progress in mid-decade and to redirect efforts in areas where we are falling behind. The groups drafting the objectives had varying approaches to setting targets, including (1) detailed demographic models of the impact of changes in risk factors and the use of preventive services on disease incidence and mortality, (2) simple and sophisticated statistical trend analyses, and (3) eyeballing the most recent data points in conjunction with educated guesswork. Because of the different methods used, some of the targets—specifically, those for coronary heart disease and unintentional injuries—look somewhat optimistic or conservative when compared to each other and to mortality trends.

The analysis in this paper focuses primarily on the mortality objectives, a small fraction of the total set of objectives. The objectives address many other health status concerns—such as the incidence or prevalence of specific diseases, disabilities, or other conditions—as well as present targets for risk factor reductions and improvements in the provision of preventive services. Because they were expressed in so many different ways, we were not able to incorporate most of the morbidity and disability measures into a common framework. We were, however, able to incorporate a single objective addressing overall activity limitation into a comprehensive measure of mortality and disability.

Demographic Framework

In order to sum up the demographic impact of achieving the mortality objectives by the year 2000, we had to translate the objectives developed by PHS into a common framework for demographic calculations. As shown in Table 2, we developed a cause-of-death list that is a compromise between the specific measures used in the objectives (and the prevention relevance that they represent) and the standard groupings from the International Classification of Diseases, Ninth Revision (ICD-9)[5] used to report vital statistics. The 1987 death rates for these causes of death serve as the base for our calculations.

Infant mortality is included in the list, even though it is not a single cause of death, because the measure appears prominently in *Healthy People 2000*[1] and is frequently cited in health policy discussions. Pneumonia and influenza are grouped in a residual category, Other Causes, for which there are no objectives or the objectives cannot be translated to appropriate death rates.

To translate the mortality measures in the objectives into a common set of cause-specific death rates, we assumed that the annual rate of change implied by comparing the baseline figure and the target applied to the change in death rates from 1987 to 2000 for standard ICD groups.

One category in Table 2, heart disease, has a residual that represents the part of the category that is not covered by the target of any specific objective, primarily because prevention efforts are thought to be significant only for the heart disease deaths due to coronary heart disease. Together the heart disease residual and the overall residual made up 10% of total deaths in 1987. For heart disease and

TABLE 1—Priority Areas for *Healthy People 2000*[1] Objectives

- Health promotion
  - Physical activity and fitness
  - Nutrition
  - Tobacco
  - Alcohol and other drugs
  - Family planning
  - Mental health and mental disorders
  - Violent and abusive behavior
  - Educational and community-based programs
- Health protection
  - Unintentional injuries
  - Occupational safety and health
  - Environmental health
  - Food and drug safety
  - Oral health
- Preventive services
  - Maternal and infant health
  - Heart disease and stroke
  - Cancer
  - Diabetes and chronic disabling conditions
  - HIV infection
  - Sexually transmitted diseases
  - Immunization and infectious diseases
  - Clinical preventive services
- Surveillance and data systems

TABLE 2—Major Causes of Death: 1987 Age-adjusted Death Rates and Alternative Rates of Change to the Year 2000[6]

		Alternative Annual Rates of Change				
Causes of Death	1987 Death Rate[a]	High Mortality Trend	Low Mortality Trend	PHS Targets	Target and No Change in Residual	Target and Residual Trend
Chronic diseases						
Heart disease	169.6	−2.23	−2.37	−2.28	−1.43	−1.78
Cancer	132.9	0.13	0.06	−0.18		
Stroke	30.3	−3.04	−5.19	−3.15		
COPD[b]	18.7	3.00	0.00	2.26		
Diabetes	9.8	0.16	−2.44	−0.79		
Cirrhosis	9.1	−3.22	−3.94	−3.15		
Injuries						
Suicide	11.7	0.50	−0.32	−0.83		
Homicide	8.6	−0.98	−3.23	−1.27		
Unintentional injuries	34.6	−2.53	−2.96	−1.25		
Infant mortality	1018.5	−2.70	−4.15	−2.78		
Other causes of death						
Pneumonia and influenza	13.1	1.64	−3.44	−1.68		
All other causes	97.0	0.26	−2.11	0.00	0.00	0.26

[a] Death rates are per 100 000 population; the infant mortality rate is given per 100 000 live births.
[b] Chronic obstructive pulmonary disease.

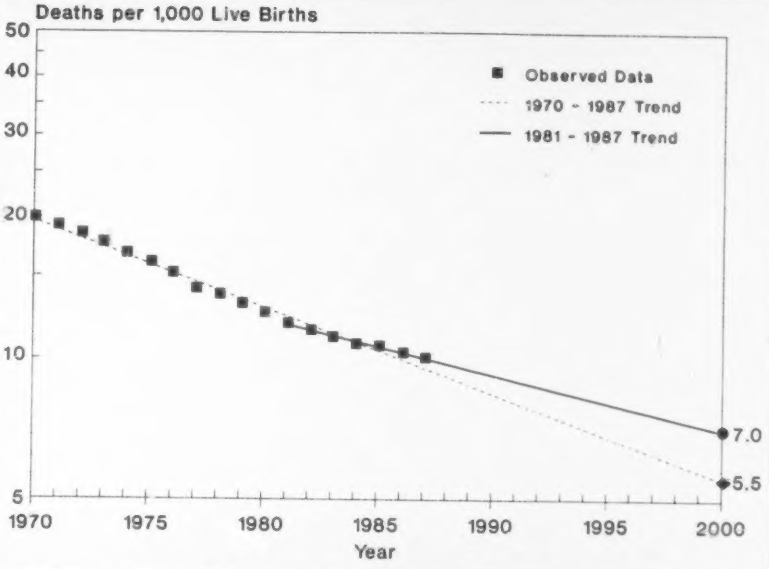

FIGURE 1—Infant mortality rate, 1970 to 1987, with 1970–1987 and 1981–1987 trends projected to the year 2000 (logarithmic scale), from *Health, United States, 1989*[4] and earlier editions of that report.

total mortality we calculated the target rate of change using three assumptions about the rate of change in the residual categories:

1. *Target:* the residual heart disease component changes at the same rate as the targeted part, and the overall residual mortality rate does not change.
2. *No change:* the residual heart disease component and the overall residual do not change.
3. *Own trend:* the residual heart disease component and the overall residual change according to the trend observed from 1979 to 1987.

Table 2 shows the three variants for the annual rate of change. The Appendix lists these and other specific assumptions made in translating the year 2000 objectives to a common demographic framework.

Trend Analyses for Comparison Purposes

To establish a reference point in a comparable manner for all of the cause-of-death categories, we chose to use simple trend analyses to estimate the potential changes in mortality that might come about even without the efforts implicit in the objectives process. More sophisticated techniques could be used here, but, given the illustrative nature of the effort, simple trend analyses are sufficient.

Because most of the rates are moving toward zero and some would become negative if linearly extrapolated to the year 2000, all of our trend analyses are in a logarithmic scale, implying a constant relative annual change. The trend analyses are based on rates for the whole population age-adjusted to the 1940 US standard million population.[4]

A look at the data going back to 1970 indicates that some trends were not simple. For example, infant mortality, as shown in Figure 1, saw a leveling off in the 1980s of the sharp downward trend experienced in the 1970s. The result is that a trend line fit to the data from 1970 to 1987 gives a very different rate of change than one fit to the data from 1981 to 1987, the years used by PHS to determine the target value for the infant mortality rate (IMR). For some other causes, the death rate is actually increasing, and the rates for other causes show substantial variability over the period.

To reflect the uncertainty in these simple trend analyses, we have developed high and low mortality variants for each cause of death. This approach assumes that the major source of uncertainty is in deciding which points to include in the trend calculation, not variation around the trend line. For most causes, the high and low variants reflect the different trend estimates calculated using data from 1970 to 1987 and data from 1979 or later to 1987. The 1970 starting point was chosen arbitrarily. The 1979 starting point corresponds to the implementation of ICD-9 in the United States, and in some cases changes between ICD-8 and ICD-9 cause major discontinuities in the data. For some causes of death PHS based the target on a trend analysis starting after 1979; for those causes we considered the additional trend line in establishing high and low mortality variants. For chronic obstructive pulmonary disease (COPD) the trend in the death rate is increasing, and we assumed a zero rate of change for the low mortality alternative. Table 2 shows the annual rate of change for the high and low mortality projections for each cause of death.

Calculation of Summary Mortality Measures

We present our results in terms of standard life table parameters e_0 (life expectancy at birth), l_{65} (the probability of surviving to age 65), and l_{65} (life expectancy at age 65). Further calculations suggest that the major qualitative conclusions regarding mortality are seen in these measures alone. The calculations were done as follows.

We started with the age- and cause-specific rates for 1987,[6] using the following age ranges: 0, 1, 5(10)85. For each cause we then applied the relative annual rate of change from Table 2 to each 1987 age-specific rate for that cause to calculate cause-specific death rates for the year 2000. For infant mortality we assumed that the targeted rate of change for IMR applied to $_1q_0$. We did not take cause-specific mortality rates into account in calculating $_1q_0$.

These cause-specific rates were aggregated into age-specific $_nM_x$ rates, and transformed into life table $_nq_x$ rates using Chiang's a's,[7] modified for 10-year intervals. Although these abridged calculations are not exact, if we assume no change for every cause, we can reconstruct the 1987 life table parameters[6] without excessive error. For instance, the actual and calculated values for e_0, e_{65}, and l_{65} are as follows:

Parameter	Actual	Calculated
e_0	75.0	75.1
l_{65}	16.9	17.2
l_{65}	79 040	79 139

Because most of the objectives are stated in terms of total population rates, we have not attempted to differentiate targets or trends in age, race, or sex-specific rates. All of the rates of change in Table 2 are for the total population.

In order to consolidate these calculations on the impact of changes in mortality with a targeted reduction in disability, we adapted methodology developed by Sullivan[8] and used by others[9,10] that allows us to estimate an expectation of life free of disability. The measure of disability that we use is drawn from Objective 17.2 in *Healthy People 2000,*[1] which calls for a decline in the proportion of the population with major activity limitations from 9.4% in 1988 to 8.0% in 2000, a decrease of 1.34% per year. The baseline data come from the National Health Interview Survey (NHIS), which provides estimates of the proportion of the civilian noninstitutionalized population that experiences a limitation in their "major activity" due to a chronic health condition.[11] Major activity is defined as the predominant social role expected of a person of a given age: playing for children under age 5, attending school for children aged 5 to 17, working or keeping house for adults aged 18 to 69, and living independently for adults aged 70 and older.

We assumed that the 1.34 annual percentage decline between the baseline and the target would apply in each of the age groups in which the NHIS data are tabulated. For the base year and 2000, we calculated the number of life table years lived with major activity limitation by multiplying the prevalence of major activity limitation in each age group by the number of life table years lived in that age group ($_nL_x$) and summing the products for age x and above. The result can be expressed as the life expectancy with major activity limitation. Subtracting that value from total life expectancy gives an estimate of expected years of life free of limitation in major activity. Because the NHIS figures exclude the institutionalized population, including residents of nursing homes and other long-term care facilities, our calculations underestimate time with activity limitations.

Results

We have calculated the effect of meeting the objectives on overall mortality and on life expectancy without limitation in major activity, examined the impact of meeting specific objectives and groups of objectives on overall mortality and on the distribution of causes of death, and carried out an in-depth exploration of the specific targets chosen for coronary heart disease and unintentional injuries.

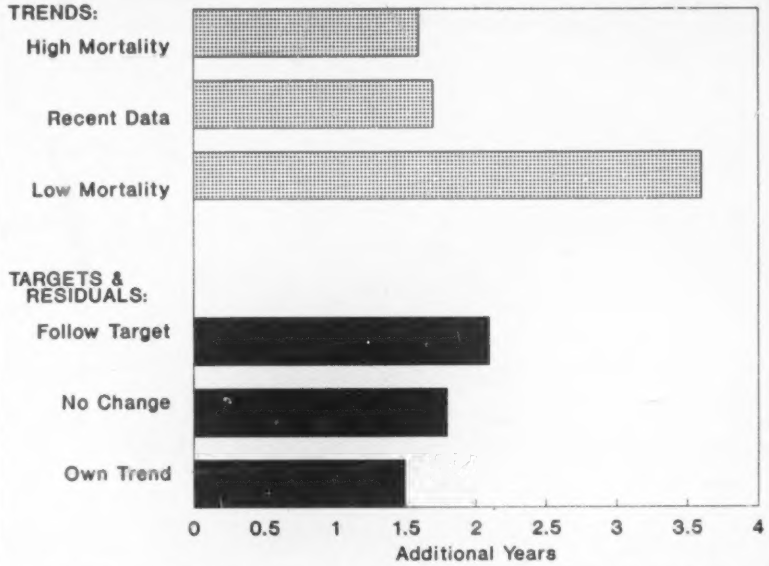

FIGURE 2—Projected increases in life expectancy at birth from trend and target changes in mortality from all causes: 1987–2000.

Overall Mortality

The effect of meeting the mortality objectives on overall life expectancy is shown in Figure 2. Depending on which of the two trend variants is used, the increase in life expectancy at birth by the year 2000 can range from 1.6 to 3.6 years. Although the trends based on data from the 1980s are not the least optimistic for every cause of death, they are in most cases. If trends based only on the more recent data were assumed, the increase in life expectancy at birth by the year 2000 would be about 1.7 years. Figure 2 also shows that the increase in life expectancy at birth to be expected if the *Healthy People 2000*[1] objectives are met is from 1.5 (assuming trends for the nontargeted causes of death) to 2.1 years (if residual rates improve at the same rate as the targeted rates). These figures imply a life expectancy of 76.6 to 77.2 years. Life expectancy at age 65 increases by 1.0 to 2.5 years with trends alone, and by 1.0 to 1.4 years if the year 2000 objectives are achieved. The current life expectancy at age 65 is 17.2 years.

Similarly, the increase in l_{65}, the life table probability of living to age 65, increases by 1.9 to 3.7 percentage points according to the two trend variants, but from 1.8 to 2.3 points if the year 2000 objectives are achieved. Thus, meeting the mortality targets would increase the probability of surviving to age 65 from 79.1% to 80.9–81.5%. Alternatively, one could say that the probability of a "premature death" would decrease from 20.9% to 18.5–19.1%. For all three measures, the targets are closer to the low end of the range defined by the alternative trends than the high end of the range.

Years without Limitations in Major Activity

With current mortality and disability statistics, an average of 66.8 years out of 75.1—89% of life expectancy at birth—would be spent free of limitation in major activity. If the target rate for major activity limitation is achieved, however, the number of years without limitation will increase to 69.3 to 69.7 years—approximately 90% of total life expectancy—depending on the assumption about trends in nontargeted causes of death as above. The scenario that produces the largest increase in total life expectancy also produces the largest increase in disability-free life. That same scenario, however, also results in the greatest remaining number of years with activity limitation (7.5 years) because relatively more people survive to the older ages at which disability is most prevalent.

Under current conditions, an average of 13.2 disability-free years would be experienced after age 65—about 77% of the life expectancy at that age. Assuming that the year 2000 objectives are met, the number of disability-free years experienced af-

TABLE 3—Estimated Impact on e_0 and l_{65} of Projected and Targeted Changes in Death Rates for Individual and Grouped Causes of Death: Year 2000

	Projected Changes		Targeted Changes		
Causes of Death	High Mortality Trend	Low Mortality Trend	PHS Targets	Target and No Change in Residual	Target and Residual Trend
Added years of life expected at birth (e_0)					
All causes	1.6	3.6	2.1	1.8	1.5
Chronic diseases	1.3	1.9	1.5	1.3	1.1
Heart disease	1.2	1.2	1.2	0.9	0.8
Cancer	−0.1	0.0	0.1		
Stroke	0.3	0.4	0.3		
COPD[a]	−0.2	0.0	−0.1		
Diabetes	0.0	0.1	0.0		
Cirrhosis	0.1	0.1	0.1		
Injuries	0.3	0.4	0.2		
Suicide	0.0	0.0	0.0		
Homicide	0.0	0.1	0.0		
Unintentional injuries	0.3	0.3	0.1		
Infant mortality	0.2	0.3	0.2		
Other causes of death	−0.2	0.7	0.1	0.1	0.0
Pneumonia and influenza	−0.1	0.1	0.1		
All other causes	−0.1	0.5	0.0	0.0	−0.1
Increase in percentage surviving to age 65 (l_{65})					
All causes	1.7	3.7	2.3	2.1	1.8
Chronic diseases	1.3	1.9	1.6	1.4	1.2
Heart disease	1.2	1.3	1.2	1.0	0.8
Cancer	−0.1	0.0	0.1		
Stroke	0.2	0.3	0.2		
COPD[a]	−0.2	0.0	−0.2		
Diabetes	0.0	0.1	0.0		
Cirrhosis	0.2	0.2	0.2		
Injuries	0.5	0.7	0.4		
Suicide	0.0	0.0	0.1		
Homicide	0.1	0.2	0.1		
Unintentional injuries	0.5	0.5	0.2		
Infant mortality	0.2	0.3	0.2		
Other causes of death	−0.2	0.8	0.0	0.0	0.0
Pneumonia and influenza	−0.1	0.1	0.0		
All other causes	−0.1	0.7	0.0	0.0	−0.1

[a] Chronic obstructive pulmonary disease.

ter age 65 increases to 14.6–15.0 years—about 80% of the life expectancy at that age.

Changes in Specific Causes of Death

To understand these results better we need to explore the contributions of specific causes and groups of causes to these summary measures. Table 3 shows the effect on e_0 of trend and target changes for each cause of death, assuming no changes in mortality for other causes. The potential changes in heart disease will have the largest impact on the increase in life expectancy (0.8 to 1.2 years), followed by stroke (0.3 to 0.4 years) and infant mortality (0.2 to 0.3 years). This should be compared to a range of 1.5 to 3.6 years of potential increase in overall e_0. Cancer mortality trends are increasing slightly, and the effect of meeting the target would change a 0.1 year decrease in life expectancy to an increase of 0.1 year. With only two exceptions (injury and "other"), none of the variants of the other causes of death lead to a change in life expectancy of more than 0.2 years.

Table 3 also shows the effect on l_{65} of trend and target changes for each cause of death, assuming no changes in mortality for other causes. The same causes as above are also the most prominent in their effect on l_{65}. Unintentional injury deaths, however, achieve more prominence.

Because of competing risks, looking at changes in one cause of death at a time tends to diminish the apparent effect of health promotion and disease prevention efforts.[12] The chronic diseases, for example, are all related to a common set of risk factors (diet, smoking, alcohol, exercise) and would all tend to benefit from any increases in the accessibility of clinical preventive services for any one of them.[13] Therefore, we decided to look at the potential for reducing mortality for all chronic diseases. Similarly, since risk factors for injury deaths would tend to run together, we will also look at intentional and unintentional injuries taken together. Infant mortality is itself a combination of causes of death at ages under 1 year, and the final category is the combination of pneumonia and influenza with the residual "other," which includes AIDS and other infectious diseases.

Figure 3 and Table 3 show the impact of trend and targeted changes on e_0 and l_{65}. The major impact in either scale is still the chronic diseases (1.1 to 1.9 additional years of life expected at birth; 1.2 to 1.9 percentage point increase in the probability of living to age 65). Because they improve survival at younger ages, changes due to reduction in infant mortality and injuries are relatively more prominent in l_{65} than in e_0, but are still less important than changes in the chronic diseases.

Another way of looking at the projected changes is to look at the changes in the distribution of the causes of death. The distribution of deaths was determined by applying the projected age- and cause-specific mortality rates to the Census Bureau's median projected population for the year 2000.[14] The ranges indicate the highest and lowest proportions resulting from the three assumptions for nontargeted rates. Based on these calculations, chronic diseases fall from 72.7% to 70.5–71.2% of all deaths; the proportion of deaths due to injuries stays about the same: 6.1% to 5.8–6.1%; infant deaths fall from 1.3% to 1.0%; and other deaths increase from 20.0% to 21.9–22.4%. Clearly, achieving the targets would not mean a major shift in the distribution of causes of death, at least in these broad categories.

For specific causes, however, there are more marked changes. The largest drop is for heart disease: 36.9% to 31.8–34.1% of deaths. The proportion also goes down for stroke: 7.5% to 5.5–5.7%. For cancer, however, the proportion increases from 21.7% to 23.6–24.6%. COPD also increases, from 3.6% of deaths to 5.3–5.6%. The smaller shares for heart disease and stroke are nearly canceled out by the larger shares for cancer and COPD, leading to little overall change in the proportion of deaths from chronic diseases.

Reevaluation of Heart Disease and Injury Targets

Further analysis shows that the targeted reductions in heart disease and unintentional injuries are less than the trends would predict. Figure 4, for instance, shows trends in coronary heart disease since 1972. (As used in *Healthy People 2000*[1] and elsewhere by PHS, coronary heart disease refers to specific ICD codes within the broader category of Diseases of the heart.[15]) Depending on which years are used for trend analysis, the projected death rate for the year 2000 is either 90.8 or 92.6 per 100 000. The target is set conservatively at 100 per 100 000, although the trend data suggest that 90 per 100 000 might be a more reasonable level.

Figure 5 and Table 4 show that changing the target to 90 per 100 000 would have a major impact on the gain in e_0 (an additional increase of about 0.3–0.4 years) and in l_{65} (an additional increase of 0.3–0.4 percentage points). It would also affect the proportion of deaths due to heart disease and to chronic diseases. Assuming 1987 rates, 36.9% of all deaths are due to diseases of the heart. The percentage would fall to 29.5–32.4% with the new target. With the original heart disease target, this drop would only be to 31.8–34.6%. Similarly, the proportion of all deaths due to chronic diseases with 1987 rates is 72.7% and would fall to 69.6–70.4%. With the original heart disease target, this drop would only be to 70.5–71.2%.

Figure 5 and Table 4 also show the impact of a change in the target for unintentional injury mortality. A report has shown that accident mortality of all types has fallen by 21% between 1979 and 1989.[16] This rate of change is about double that called for in the PHS objectives. Changing the target from 29.3 to 22.0 per 100 000, consistent with the 10-year trend, would lead to a 0.2–0.3 year increase in life expectancy at birth and a 0.3–0.4 percentage point increase in the probability of surviving to age 65. The proportion of deaths due to all injuries would drop from 6.1% to 4.9–5.1% of all deaths.

Taken together, lowering both the heart disease and unintentional injury targets would increase overall life expectancy at birth by 0.6 years to 77.2–77.8 years. The probability of surviving to age 65 would increase by 0.6–0.8 percentage points.

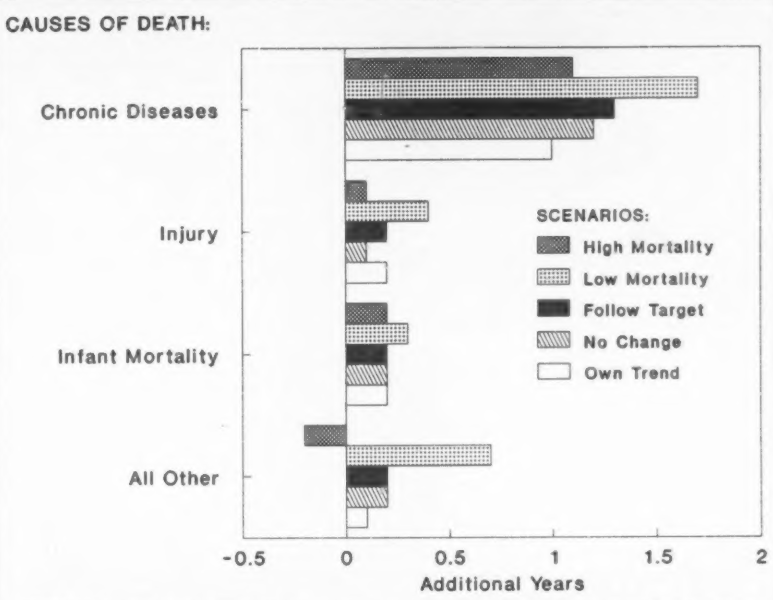

FIGURE 3—Projected increases in life expectancy at birth from trend and target changes in mortality from specified groups of causes: 1987–2000.

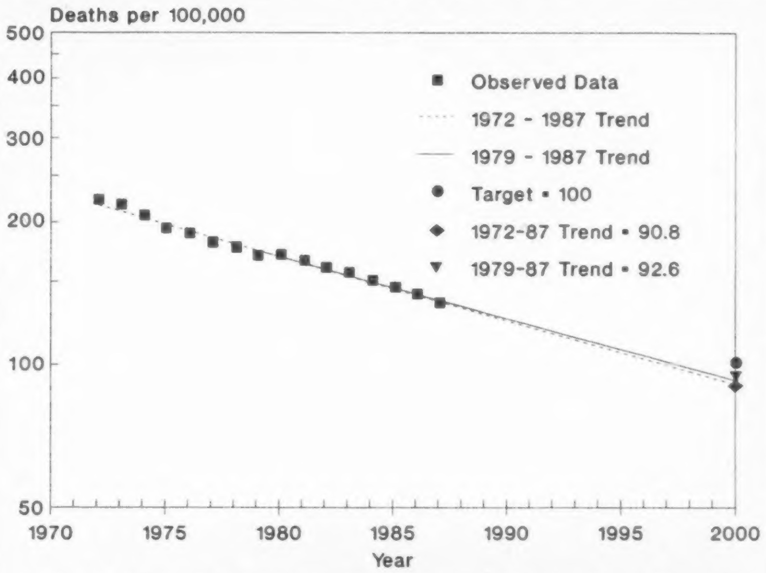

FIGURE 4—Age-adjusted coronary heart disease death rate, 1972 to 1987, with 1972–1987 and 1981–1987 trends projected to the year 2000, and year 2000 target rate (logarithmic scale, 1940 standard population). Data for 1972 to 1987 calculated by C. Sempos, National Center for Health Statistics, and provided by the Office of Disease Prevention and Health Promotion, US Public Health Service.

Discussion

Meeting the mortality targets set in *Healthy People 2000*[1] would have an important demographic impact. Taking the coronary heart disease and unintentional injury targets as stated in the objectives, the increase in life expectancy at birth would be 1.5 to 2.1 years, implying a life expectancy of 76.6 to 77.2 years. If the coronary heart disease and unintentional injury targets were changed to be somewhat more optimistic than the recent trend, life expectancy at birth would increase by 2.1 to 2.7 years to 77.2 to 77.8 years (see Figure 6). Trends in the cause-specific death rates used in this paper suggest an increase of 1.6–3.6 years to 76.7–78.7 years, with the lower end of the range being more likely. The results can also be

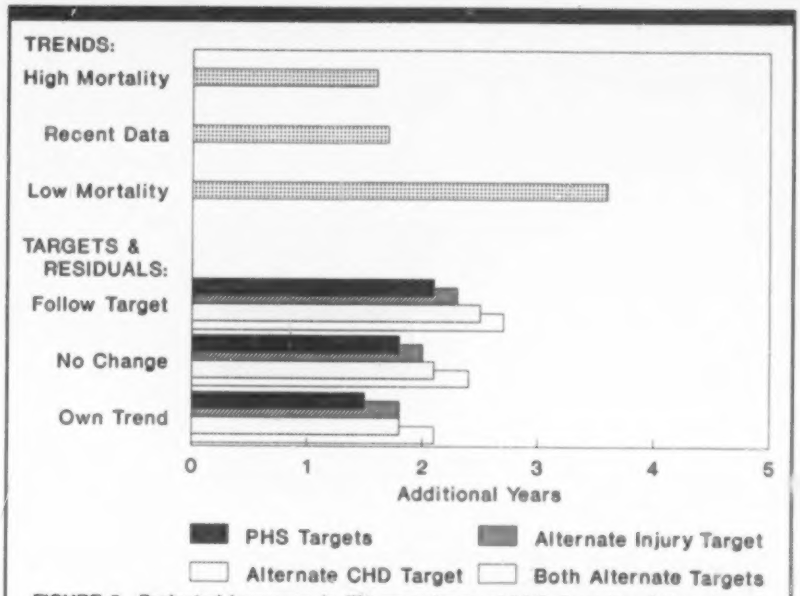

FIGURE 5—Projected increases in life expectancy at birth from trend, target, and alternative target changes in mortality from all causes: 1987–2000.

TABLE 4—Estimated Impact on e_0 and l_{65} of Alternative Targets for Coronary Heart Disease and Unintentional Injuries and Their Associated Grouped Causes of Death: Year 2000

	Target Changes		
Causes of Death	Alternate Targets	Target and No Change in Residual	Target and Residual Trend
Added years of life expected at birth (e_0)			
All causes	2.7	2.4	2.1
Heart disease only	2.5	2.1	1.8
Unintentional injuries only	2.3	2.0	1.8
Chronic diseases	1.9	1.6	1.4
Heart disease	1.6	1.2	1.0
Injuries	0.4		
Unintentional injuries	0.3		
Increase in percentage surviving to age 65 (l_{65})			
All causes	3.1	2.7	2.4
Heart disease ony	2.7	2.4	2.1
Unintentional injuries only	2.7	2.4	2.2
Chronic diseases	2.0	1.7	1.5
Heart disease	1.6	1.3	1.1
Injuries	0.7		
Unintentional injuries	0.6		

compared to ranges used in current Census Bureau[14] and Social Security Administration[17] projections of 75.3 to 78.9 years and 75.9 to 78.6 years, respectively. These are based on extrapolations as well, but using a different methodology. As Figure 6 shows, meeting the targets with the changes in the coronary heart disease and unintentional injury targets suggested above would put overall life expectancy above the middle of these ranges.

There are two reasons why demographic forecasters ought to note these results. First, they represent the substantive input of public health professionals asked to think about what might be achievable with known interventions. This kind of input is something that demographers would like to have but often cannot obtain. Second, the public health establishment is committed to making these changes come about. If progress reviews in mid-decade indicate that the nation is falling behind in measures needed to meet these targets, more resources might actually be brought to bear on interventions to reduce mortality, closing the feedback loop.

It is important to note, however, that the process used to set the targets might not have used the substantive knowledge of the experts as effectively as possible. Much of the energy used in setting the objectives has gone into discussions about which subjects should be included in the list of priority areas, into determining specific issues to be addressed within those areas, and into deciding how the objectives should be expressed in quantitative terms. In some cases, the specific numerical targets emerged after careful analyses and thought, but in others they were set at the last minute with little technical or substantive input. The process might have been improved if the content and the measures of the objectives had been determined first, and then a concerted effort made across the board to set target levels based on trend analyses and mathematical models that were feasible, realistic, and internally consistent.

Meeting the objective for reductions in activity limitation would lead to an additional 2.5 to 2.9 years of limitation-free life at birth and 1.4 to 1.8 years of limitation-free life after age 65. Chronic conditions of the circulatory system and injuries are major causes of disability as well as mortality.[18] Depending on whether increased survival consistent with the targets suggested in this article comes about through primary prevention or through better medical treatment of persons with heart disease and injury, further improvements in limitation-free years may be possible. Our analysis, however, focuses on mortality rather than other important health status variables such as morbidity and disability, and the available information does not permit explicit calculations.

The analysis does not specifically study the effect of projected changes in risk factors and preventive services. Nor does it address the question of different age patterns of change in mortality rates. The analysis needs to be expanded in order to appreciate the full range of benefits to be achieved and to better understand the contributions that various interventions can make to improvements in health status. As discussed above, this has not been possible for two reasons: lack of a consistent framework for nonmortality health status measures and lack of complete information on the effect on health status of changes in risk factors and use of preventive services.

Progress in expanding the discussion beyond mortality alone depends on the development of consistent, systematic, and timely measures of health status, risk factors, and preventive services. Data from the annual NHIS are used for many of the objectives, and with changes in the objectives and the survey a more integrated system might be achievable.

Erickson and her colleagues have suggested that composite health status measures that incorporate morbidity, disability, and quality of life could be obtained from the NHIS, and that these could be the basis for the calculation of a measure of quality-adjusted life years (QALY) that is analogous to life expectancy except that each year of life is discounted by the average "quality of life" people experience at that age.[19] To the extent that the year 2000 objectives for the prevalence of specific conditions relate to the measures that go into the QALY calculation, the impact of meeting those objectives could be summed up in the same way we have done here for objectives on mortality levels.

Because the PHS objectives are generally stated in terms of overall death rates without specific mention of age or sex patterns, we assumed for this analysis that the same percentage change applied to all age-specific death rates and did not differentiate by sex. The current patterns of risk reduction and use of preventive services probably do not result in the same percentage reductions in death rates at all ages, as we were forced to assume. Clearly, further analysis needs to be used using more complex models of mortality decline than the ones we have used. Manton and his colleagues[20] and Olshansky,[21] for instance, have investigated the impact of delaying the progression of chronic diseases, not reducing their incidence.

Making the link between changes in risk factors and preventive services to health status measures will require two things. First, we need more extensive epidemiological studies on the prevalence of risk factors and the effect of changes in risk factors and preventive services on a full range of health status measures. Furthermore, preventive interventions must be carefully evaluated so that their population impact can be better understood. Second, more extensive demographic and epidemiological modeling efforts are needed to bring this information together so that the impact of changes in risk factors and access to clinical preventive services can be better predicted. There is some work in this area that could be drawn on,[22-27] but generally for one disease or risk factor at a time, and it is difficult to integrate this information.

Models of this sort can also shed light on the age-specific impact of preventive interventions that is lacking in this analysis. Manton, for instance, argues that risk factors and mortality reductions have a very different relationship at older ages than they do at younger ages.[28] Two kinds of issues need to be explored. First, detailed studies of the actual patterns of mortality reductions would lead to a fuller understanding of the demographic impact of meeting the year 2000 objectives. Second, we need to explore the impact of interventions targeted at different age groups and thus get a better idea about the most effective way to manage health promotion and disease prevention interventions. □

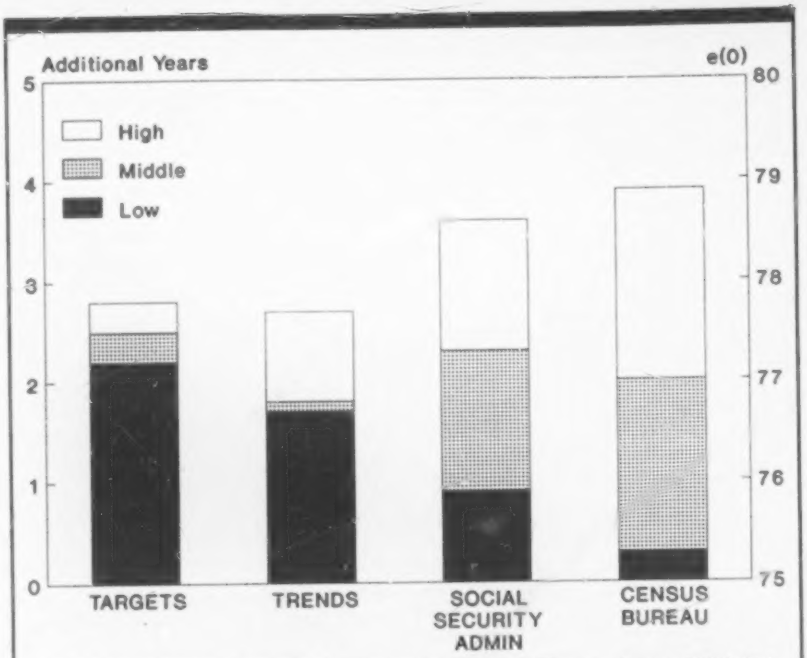

FIGURE 6—Alternative ranges of projected increases in life expectancy at birth (e_0) from 1987 to 2000 and of projected values of e_0 in the year 2000: PHS targets, trends, Social Security Administration, and Bureau of the Census. Census Bureau values are arithmetic averages of separate projections for males and females. From Spencer[14] and Wade[17].

Acknowledgments

This work was supported by the Office of Disease Prevention and Health Promotion, Office of the Assistant Secretary for Health, Department of Health and Human Services, under corporate agreement no. HPV-87-002-03-0. It was conducted through the Division of Health Promotion and Disease Prevention, Institute of Medicine, National Academy of Sciences.

A previous version of this paper was presented on May 4, 1990, at the annual meeting of the Population Association of America, held in Toronto, Ontario. We are grateful for comments and suggestions offered by the discussant, other panel members, and other colleagues.

The analysis and opinions in this paper are those of the authors. They do not necessarily reflect those of the Institute of Medicine, the National Academy of Sciences, or the US Public Health Service.

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APPENDIX—Demographic Assumptions

Cause-of-Death Categories (ICD-9 Codes)	*Healthy People 2000* Objectives	1987 Baseline Rate[a]	2000 Target Rate[a]	High Mortality Trend	Low Mortality Trend	Notes on Trend Data[b]
Heart disease (Diseases of the heart: 390-398, 402, 404-429)	Objective 15.1: Coronary heart disease (410-414, 402, 429.2)	135.0	100.0	1979–1987	1970–1987	Trends calculated from vital statistics rates for "Diseases of the heart"; nontargeted portion of heart disease (the residual) was calculated by subtracting the coronary heart disease rate from the total heart disease rate; coronary heart disease rates for 1972–1987 were calculated by C. Sempos, National Center for Health Statistics, and provided by the Office of Disease Prevention and Health Promotion (ODPHP), US Public Health Service.
Cancer (Malignant neoplasms: 140-208)	Objective 16.1: Cancer (all sites)	133.0	130.0	1970–1987	1979–1987	
Stroke (Cerebrovascular disease: 430-438)	Objective 15.2: Stroke	30.3	20.0	1982–1987	1972–1987	Trends calculated from rates prepared by the National Heart, Lung, and Blood Institute and provided by the ODPHP. These rates include a small adjustment in published vital statistics rates to establish consistency of ICD-8 and ICD-9 rates.
COPD (Chronic obstructive pulmonary diseases and allied conditions: (490-496)	Objective 3.3: COPD	18.7	25.0	1979–1987	—	Trend was calculated only for vital statistics rates based on ICD-9; rates under ICD-8 are not comparable.
Diabetes (Diabetes mellitus: 250)	Objective 17.9: Diabetes-related deaths (as underlying or contributing cause)	38.0[c]	34.0	1979–1986	1970–1987	High mortality trend was calculated from rates for diabetes as underlying or contributing cause of death; rates were prepared by the Centers for Disease Control (CDC) and provided by the ODPHP. Low mortality trend was calculated from vital statistics rates for diabetes as an underlying cause of death.
Cirrhosis (Chronic liver disease and cirrhosis: 571)	Objective 4.2: Cirrhosis	9.1	6.0	1970–1987	1979–1987	
Suicide (E950-E959)	Objective 6.1: Suicide	11.7	10.5	1980–1987	1970–1987	High mortality trend calculated from rates for suicide prepared by CDC and provided by the ODPHP; low mortality trend calculated from vital statistics rates for suicide.
Homicide (Homicide and legal intervention: E960-E978)	Objective 7.1: Homicide (E960-E969)	8.5	7.2	1970–1987	1960–1987	High mortality trend calculated from vital statistics rates for "Homicide and legal intervention"; low mortality trend calculated from rates for "homicide" alone; rates were prepared by CDC and provided by the ODPHP.
Unintentional injuries (Accidents and adverse effects: E810-E949)	Objective 9.1: Unintentional injuries	34.5	29.3	1960–1987	1979–1987	High mortality trend calculated from rates for unintentional injuries prepared by CDC and provided by the ODPHP; low mortality trend calculated from vital statistics rates for "Accidents and adverse effects."
Infant mortality (all deaths at less than 1 year old)	Objective 14.1: Infant mortality	10.1	7.0	1981–1987	1970–1987	Rates are from *Health, United States, 1989*[4] and earlier editions.
Pneumonia and influenza (480-487)	Objective 20.2: Epidemic-related pneumonia and influenza deaths among people ages 65 and older	9.1	7.3	1979–1987	1970–1987	
All other causes	—	—	—	1979–1987	1970–1987	Rates for each year in this category determined by subtracting vital statistics rates for the specified causes of death from the total death rate for the year.

[a] Source: Reference 1. Rates, except for the infant mortality rate, are per 100 000 population and are age-adjusted to the 1940 standard population. The infant mortality rate is per 1000 live births.

[b] Source: Reference 3, except where noted.

[c] Baseline is for 1986.

Public Health Briefs

Injury Prevention Awareness in an Urban Native American Population

James S. J. Hsu, BS, and Scott D. Williams, MD, MPH

ABSTRACT

The injury-related mortality rate for Native American children between the ages of 1 and 4 years is nearly three times that of the same age group in the general population. To assess injury prevention awareness in urban Native American families, we administered 39 age-appropriate questions from the Framingham Safety Survey to 50 Native American families and 100 other families and developed an answer scoring system to analyze and compare survey responses. Survey responses revealed that Native American families are less likely to keep small objects, household products, and medicines out of the reach of their children and to possess and understand the use of ipecac. Although urban Native-American families appear to be less aware of ingestion prevention practices than other urban families, these and other deficiencies in injury prevention awareness are more likely the result of factors related to their low-income status than to culturally based practices. (*Am J Public Health.* 1991;81:1466–1468)

Introduction

Although nearly half of all Native Americans now live in urban areas, little data are available describing the health status for this segment of the population. Health statistics reported by the Indian Health Service often describe only reservation-dwelling Native Americans. Among Native American children between 1 and 4 years of age, injury-related deaths occur at nearly three times the rate of the same age group among the general population in the United States.[1–4] As with most accident-related injuries and deaths, a large portion of those involving Native American children may be preventable. Urban conditions such as poor housing, limited access to health care, inadequate day care, and increased exposure to physical hazards may further increase the risk of injury for urban Native American children.

Injury prevention awareness is commonly evaluated in the clinical setting by parent survey techniques such as the Framingham Safety Survey (FSS), recommended and distributed by the American Academy of Pediatrics (AAP) through The Injury Prevention Program (TIPP).[5] These survey methods have never been studied for their usefulness in identifying injury prevention awareness in the Native American population. In this study we attempted to determine whether parents in an urban Native American population are as aware of, and as likely to practice, injury prevention techniques as parents in other urban families.

Methods

In this cross-sectional study, we enrolled 150 families with children between the ages of 1 and 4 during the 4-month period from June 1988 to September 1988. Fifty families were interviewed at each of three clinics in the same urban area: (1) the Indian Health Care Clinic, a primary health care clinic for Native American families; (2) the Salt Lake City–County Health Department Clinic, a primary health care clinic for low-income families; and (3) a university-based private pediatric practice.

To assess injury prevention awareness in these families, a trained interviewer verbally administered 39 age-appropriate questions, taken from the FSS, to a parent from each enrolled family while the families were waiting at the clinic for a well-child or sick-visit appointment. A scoring method was devised to quantify the risk of injury. A score of 1 was given for low-risk, 2 for moderate-risk, and 3 for high-risk behavior. A total score was computed for each family, the lower total scores being associated with a lower risk of injury. The results were analyzed for statistical significance using the Student's *t* test, the Kruskal-Wallis test, and the Mann-Whitney test. Statistical significance was determined at the $P < .05$ level.

Results

Characteristics of the families in the sample are presented in Table 1. There was a significant difference in income between families at the three clinics ($P = .001$), with families at the urban Indian clinic having a lower mean annual income than families at either of the other

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The authors are with the University of Utah School of Medicine.

Requests for reprints should be sent to Scott D. Williams, MD, University of Utah School of Medicine, Department of Pediatrics, 50 North Medical Drive, Salt Lake City, UT 84132.

This paper was submitted to the journal August 20, 1990, and accepted with revisions April 24, 1991.

two clinics. When families were stratified into a low-income (<$15 000/year) group and a higher income (>$15 000/year) group, the Native American families within the low-income group had a significantly lower mean annual income than other families in the low-income group (*P* = .01). Further income stratification was not performed because of limitations of the sample size.

Table 2 presents the item scores on questions where statistically significant differences existed between groups. Of the 39 injury prevention awareness questions, Native American families received higher risk scores on questions 1 through 4, even after controlling for income. Low-income families, in general, received higher risk scores on questions 1, 2, and 5 through 8. Higher income families received higher risk scores on questions 9 through 11.

Total scores are also presented in Table 2. A Native American revised score was computed, eliminating questions 1 through 4, to look for a cumulative injury prevention awareness deficiency among Native American families beyond the questions where statistically higher scores were found. To create a similar measure among all low-income families, a low-income revised score was computed, eliminating questions 1, 2, and 5 through 11.

TABLE 1—Characteristics of the Injury Prevention Awareness Study Population

	Income < $15 000/y		All Races	
	Native Americans	Non-Native Americans	Income < $15 000/y	Income > $15 000/y
Number of families	42	45	87	63
Household income/y				
$0–5000	25	19	44	
$5001–10 000	15	11	26	
$10 001–15 000	2	15	17	
$15 001–20 000				14
$20 001–25 000				10
> $25 000				39
Children aged 1–4 y				
Total children	61	67	128	78
Children/family	1.45	1.49	1.47	1.23[a]
Previous accidents resulting in:				
Visit to a clinic	3	14	17	18
Visit to an emergency room	8	16	24	16
Admission to a hospital	1	4	5	2

[a] *P* = .03.

Discussion

After controlling for income, our data revealed four items on which low-income Native American families received signif-

TABLE 2—Selected Mean Item Scores, Total Scores, and Revised Total Scores on the Modified Framingham Safety Survey Stratified by Income and Race

	Income < $15 000/y			All Races		
Questionnaire Item	Native Americans	Non-Native Americans	*P* Value	Income < $15 000/y	Income > $15 000/y	*P* Value
1. Do you have ipecac in the home?	2.81	2.04	.0001	2.41	1.67	.0001
2. Do you know how to use ipecac?	2.76	2.07	.0004	2.40	1.67	.0001
3. Do you keep plastic wrappers, bags, balloons, peanuts, and other small objects out of the reach of your children?	1.35	1.06	.02	1.21	1.21	NS
4. Do you keep household products, medicines (including baby powder, aspirin, and iron), and sharp objects out of reach and in locked cabinets?	1.10	1.0	.04	1.05	1.10	NS
5. How frequently do you check the heating system in your home?	1.69	1.56	NS	1.62	1.16	.0003
6. Do you have working fire extinguishers in the house?	2.52	2.51	NS	2.52	2.08	.0051
7. Do you have smoke or fire detectors in the house?	1.71	1.67	NS	1.69	1.25	.0025
8. What type of car seat do you use when your children ride in the car?	2.02	1.84	NS	1.93	1.64	.0132
9. Do you have a screen for the fireplace?	1.14	1.24	NS	1.19	1.51	.0004
10. Are any of your babysitters less than 13 years old?	1.19	1.33	NS	1.26	1.57	.02
11. Do you keep guns or air rifles in your house?	1.04	1.17	NS	1.12	1.67	.0001
Total score (all 39 questions)	59.10	56.70	NS	57.86	54.95	.0047
Native American revised total score[a]	51.07	50.53	NS			
Low-income revised total score[b]				41.75	40.72	NS

Note. 1 = low risk, 2 = moderate risk, 3 = high risk. NS = not significant (*P* > .05).

[a] Native American revised total score = total score − questions 1–4.

[b] Low-income revised total score = total score − questions 1, 2, and 5–11.

icantly higher scores than other low-income families. These four items are all related to ingestion events and reveal a situation where families are lacking understanding of both prevention and intervention strategies in this specific area. Because of the significant difference in income between Native American and other families, a difference that persisted even within the low-income stratum, we cannot be sure that even these four indicators of relative deficiencies in injury prevention awareness would not have been eliminated with a larger sample size including more "very-low-income" controls.

Poverty alone is known to be associated with an increase in injury-related deaths.[6-8] When compared with the higher income group, low-income Native American families displayed other deficiencies in injury prevention awareness that were common to all low-income families when compared with the higher income group. Although income bias may exist in some of the questions, the questions still reveal situations where lack of money or lack of control over housing conditions leaves a family unable to comply with injury prevention strategies such as regularly checking the heating system, using smoke detectors, and possessing a fire extinguisher. The lack of significant differences in low-income revised scores leads us to conclude that, beyond the differences revealed by individual items, no cumulative difference exists in injury prevention awareness between the low-income and higher income groups.

Care must be taken when interpreting the responses to certain items. For example, the question revealing less awareness of car seat type among low-income families may reveal a higher use of borrowed or handed-down car seats in these families, or alternatively may actually suggest a lack of confirmation regarding the family's possession and use of a car seat. Furthermore, in many instances, the "correct" response is well understood but may not actually be practiced. This suggests the possibility of obtaining falsely reassuring responses. Therefore, while responses indicating a lack of injury prevention awareness may provide guidance to the clinician in focusing health education strategies with a particular family, "correct" responses may not be enough to allow the clinician to make presumptions regarding the family's compliance.

We have demonstrated that, in families using an urban Native American clinic, there is less awareness of ingestion-related injury prevention and intervention strategies than in families using other urban clinics, although our study population was not large enough to completely eliminate income as an explanation for this difference. As members of the larger low-income urban population, urban Native American families revealed other deficiencies in injury prevention awareness or practice when compared with higher income families. Factors contributing to the increased risk of injury-related morbidity and mortality among Native American children seem to be more strongly associated with economic conditions than culturally based differences in parenting in the urban families we studied. ☐

Acknowledgments

This paper was presented at the 1989 National Student Research Forum and the 117th Annual Meeting of the American Public Health Association.

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AIDS Surveillance among American Indians and Alaska Natives

Russ Metler, RN, MSPH, George A. Conway, MD, MPH, and Jeanette Stehr-Green, MD

ABSTRACT

To assess the human immunodeficiency virus epidemic among American Indians and Alaska Natives (AI/AN), we examined acquired immunodeficiency syndrome (AIDS) case and seroprevalence data through December 1990. While AI/AN had a low 1990 reported AIDS case rate (4.0/100 000), the increase in diagnosed cases adjusted for reporting delays from 1989 to 1990 was higher (23.1%) among AI/AN than any other racial/ethnic group. Seroprevalence data for military applicants have documented higher rates for AI/AN than for either Whites or Asian/Pacific Islanders. (*Am J Public Health.* 1991;81:1469–1471)

Introduction

Although the population of American Indians and Alaska Natives (AI/AN) is small (less than 2 million), it includes highly diverse cultures and geography. Compared with the total US population, AI/AN are younger, less educated, less likely to be employed, and poorer.[1] These factors, combined with high rates of sexually transmitted diseases[2–5] and drug use among AI/AN,[6,7] may favor the spread of the human immunodeficiency virus (HIV).[8] We compared selected demographic and risk characteristics among AI/AN with the acquired immunodeficiency syndrome (AIDS) and persons with AIDS of other racial/ethnic groups reported to the Centers for Disease Control (CDC) through 1990.

Methods

State and local health departments in the United States report AIDS cases without personal identifiers to CDC. Among the variables collected are risk factors for HIV infection, setting where first AIDS diagnosis was made, and race/ethnicity (White, Black, Hispanic, Asian/Pacific Islander [A/PI], and AI/AN). Race/ethnicity is generally determined from available medical records.

AIDS cases reported to CDC as of December 31, 1990, were compared by race/ethnicity. Population data for rate calculations were based on 1990 census estimates.[9]

For AI/AN cases, the hospital of AIDS diagnosis was compared for type of organization in the *American Hospital Association Guide to the Health Care Field.*[10] Nonhospital facilities were coded consistent with the American Hospital Association's classification scheme.

All civilian applicants for US military service are tested for HIV infection.[11,12] From data provided by the Office of the Assistant Secretary of Defense (Health Affairs), HIV results for military recruit applicants examined from October 1985 through December 1990 were analyzed by race/ethnicity. Of the 2 946 895 applicants screened, 20 036 (0.7%) were AI/AN.

Results

AI/AN had 237 AIDS cases reported through 1990, the lowest cumulative total of any race/ethnicity. The rate of reported AIDS cases per 100 000 population for 1990 was low for AI/AN (4.0) relative to Blacks (42.5), Hispanics (31.9), and Whites (11.8). The rate for AI/AN was similar to that for A/PI (3.8). Comparisons based on cases diagnosed in 1989 and 1990, with adjustments made for reporting delay, showed AI/AN with the greatest relative increase (23.1%), compared with Hispanics (13.3%), Blacks (12.0%), Whites (2.5%) and A/PI (−8.8%).[13]

The exposure category distribution of AI/AN AIDS cases differed from other groups, with a smaller proportion of cases attributed to homosexual/bisexual contact (53%) than Whites (76%) but a higher proportion than for Blacks (35%) or Hispanics (39%) (Table 1). The proportion of cases attributed to intravenous drug use among women and heterosexual men in AI/AN (16%) was greater than for Whites (8%), but less than for Blacks (38%) and Hispanics (39%). AI/AN had the lowest male-to-female ratio of AIDS cases due to intravenous drug use. Of 37 AI/AN women with AIDS, 17 were intravenous drug users and 5 were heterosexual contacts of intravenous drug users. Overall, 79 (33%) of the AI/AN AIDS cases had risk factors associated with intravenous drug use.

Nearly half the AI/AN with AIDS were first diagnosed at either a federal or nonfederal government facility (Table 2). However, only 9% of all cases were initially diagnosed at an Indian Health Service (IHS) facility. In four states (Alaska, Arizona, New Mexico, and Oklahoma) with a high percentage of AI/AN served by IHS (data provided by the Division of

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The authors are with the Division of HIV/AIDS, Center for Infectious Diseases, Centers for Disease Control, Atlanta, Ga.

Requests for reprints should be sent to Russ Metler, RN, MSPH, Division of HIV/AIDS, E-47, CID, CDC, Atlanta, GA 30333.

This paper was submitted to the journal December 11, 1990, and accepted with revisions July 9, 1991.

TABLE 1—Distribution of AIDS Cases Reported through December 31, 1990, by Selected Characteristics and Racial/Ethnic Group

	Racial/Ethnic Group				
Characteristic	American Indian/ Alaska Native	White	Black	Hispanic	Asian/ Pacific Islander
No. of cases	237	88 339	45 460	25 638	996
Male/female ratio					
Among all cases	5.4:1	19.3:1	4.2:1	6.3:1	10.4:1
Among IVDU[a]	1.3:1	3.2:1	2.8:1	5.2:1	2.1:1
Exposure group					
Male homosexual/bisexual contact and IVDU (%)	13	7	6	6	2
Male homosexual/bisexual contact (%)	53	76	35	39	74
Heterosexual IVDU (%)	16	8	38	39	4
Heterosexual contact (%)[b]	3	1	5	5	2
Other (%)	15	8	16	11	18

Note. IVDU = intravenous drug use.

[a] IVDU among women and heterosexual men only.

[b] Heterosexual contact with a person with or at risk for human immunodeficiency virus infection.

TABLE 2—Affiliation of Facility Where the Diagnosis of AIDS Was Established for American Indians/Alaska Natives

Affiliation	No.	%
Nonfederal government	83	35
Federal government	33	14
Indian Health Service	22	9
Veterans Administration and other federal	11	5
Nongovernment	109	46
Church	26	11
Other	83	35
Not categorized[a]	12	5
Total	237	100

[a] Not categorized because of insufficient information provided on the case report.

Program Statistics, IHS, January 1988), 23 of 53 (43%) AI/AN AIDS cases were IHS diagnosed. In New York and California, where most AI/AN individuals are not served by IHS, none of the 75 AI/AN AIDS cases were initially diagnosed by IHS.

Among civilian applicants for military service, 17 AI/AN were HIV-antibody positive. Seroprevalence rates per 1000 were higher for AI/AN (0.8) than for Whites (0.5) and A/PI (0.4) but lower than those for Blacks (3.3) and Hispanics (1.8).

Discussion

The HIV epidemic among AI/AN may be more extensive than previously recognized. AI/AN had the largest percentage increase in diagnosed cases from 1989 to 1990 among all racial/ethnic groups. Data from applicants for military service and reports on racial misclassification, AI/AN drug use, and sexually transmitted disease rates support this possibility.

Among civilian applicants to military service, AI/AN HIV seroprevalence rates are higher than those for Whites. This higher rate relative to AIDS case rates might be the result of (1) the possibility that military recruit applicant screening data do not accurately represent the HIV infection rate for each racial/ethnic group, (2) HIV having only recently entered the AI/AN population, and (3) race ascertainment possibly differing for AIDS cases and applicants.

The incidence of AIDS among AI/AN may be higher than recorded, especially in nonreservation areas, because of racial misclassification. Several studies have documented racial misclassification in health reports,[14–16] and misclassification of AI/AN AIDS cases has been found in two cities (M. G. Hurlich, written communication, June 1990).[17]

Studies comparing drug use among AI/AN youth and non-AI/AN youth have revealed higher rates of drug use among AI/AN than non-AI/AN.[6,7] Overall, one third of all AI/AN cases were in some way associated with intravenous drug use.

High sexually transmitted disease rates could be a factor in the HIV epidemic among AI/AN. The presence of sexually transmitted diseases may facilitate HIV transmission[18–21] and reflect high-risk sexual activity. From 1984 to 1988, in the 13 states where 59% of the nation's AI/AN population reside, the mean rates of gonorrhea and primary and secondary syphilis in AI/AN were more than twice those for non-AI/AN.[2]

IHS facilities provide services for approximately 61% of the entire AI/AN population;[2] however, only 22 (9%) AI/AN AIDS cases were initially diagnosed at these facilities. This might reflect lower availability of IHS services in geographic areas with higher numbers of AI/AN AIDS cases.

Although the rate of reported AIDS cases in AI/AN has been relatively low, AIDS surveillance and HIV seroprevalence data clearly show that HIV has entered the AI/AN population, with diagnosed AI/AN AIDS cases increasing more rapidly than in any other racial/ethnic group. In addition, disturbingly high sexually transmitted disease and drug use rates have been reported among AI/AN. Culturally sensitive prevention programs for AIDS, sexually transmitted diseases, and intravenous drug use must be refined and expanded to decrease further spread of HIV among American Indians and Alaska Natives. ☐

Acknowledgments

This paper was presented in part at the 6th International Conference on AIDS in San Francisco, June 21, 1990.

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Hearing Loss and Hearing Aid Use in Hispanic Adults: Results from the Hispanic Health and Nutrition Examination Survey

David J. Lee, PhD, Deborah L. Carlson, PhD, Heidi M. Lee, MA, Laura A. Ray, MA, and Kyriakos S. Markides, PhD

ABSTRACT

Data from the Hispanic Health and Nutrition Examination Survey were employed to investigate the prevalence of hearing loss and hearing aid use in Mexican-American, Cuban-American, and Puerto Rican adults. Hearing loss was 6 to 14 times more prevalent in older (ages 54 to 74) vs younger (ages 20 through 34) subjects. Cuban Americans and Mexican Americans tended to have a similar prevalence of hearing loss, whereas Puerto Ricans had markedly lower rates. Mexican-American men had higher rates of hearing loss than Mexican-American women. The prevalence of hearing aid use among hearing-impaired individuals ranged from 2% to 11%. Implications for future research are discussed. (*Am J Public Health.* 1991;81:1471–1474)

Introduction

An estimated 28 million people in the United States are afflicted with hearing loss.[1] Hearing impairment is the fifth most prevalent chronic health condition and the second most prevalent impairment[2] in the nation. The prevalence of hearing loss is significantly greater in the elderly and also in men.[1,3] African Americans have lower rates of hearing loss than non-Hispanic Whites,[4–6] possibly because greater cochlear pigmentation may reduce the risk of noise-induced hearing loss.[7]

There have been several national population-based audiometric surveys of non-Hispanic Whites and African Americans,[4–6] but only recently has a similar survey been undertaken on the US Hispanic population. This report provides prevalence estimates of hearing loss and hearing aid use in Hispanics using data from the Hispanic Health and Nutrition Examination Survey (HHANES).

David J. Lee is with the Department of Epidemiology and Public Health, University of Miami School of Medicine. Deborah L. Carlson is with the Department of Otolaryngology, and Laura A. Ray and Kyriakos S. Markides are with the Department of Preventive Medicine and Community Health, University of Texas Medical Branch, Galveston, Tex. Heidi M. Lee is with the Department of Speech Language Pathology and Audiology, Miami Children's Hospital, Miami, Fla.

Requests for reprints should be sent to David J. Lee, PhD, University of Miami School of Medicine, Department of Epidemiology and Public Health, (R-669), PO Box 016069, Miami, FL 33101.

This paper was submitted to the journal October 30, 1990, and accepted with revisions July 9, 1991.

TABLE 1—Prevalence of Hearing Loss by Age Group

	Age Group						
	20–34		35–54		55–74		
	%	SE	%	SE	%	SE	χ^2
Mexican Americans							
Men	7.4	.02	9.4	.02	46.3	.03	38.8*
Women	2.6	.01	7.6	.02	34.9	.02	83.0*
Cuban Americans							
Men	4.6	.05	13.0	.04	48.1	.05	26.6*
Women	3.2	.04	10.9	.03	42.7	.04	25.1*
Puerto Ricans							
Men	2.3	.03	10.3	.03	23.1	.05	11.4*
Women	2.8	.02	5.5	.02	33.2	.04	23.1*

Note. Logistic regression techniques were used to test age group differences.
*$P < .001$.

TABLE 2—Prevalence of Mild (26–40 dB) and Moderate or Greater (> 40 dB) Hearing Loss by Age Group

	Mild							Moderate						
	20–34		35–54		55–74			20–34		35–54		55–74		
	%	SE	%	SE	%	SE	χ^2	%	SE	%	SE	%	SE	χ^2
Mexican Americans														
Men	4.6	.01	6.5	.02	31.7	.03	27.8***	2.9	.01	2.9	.01	14.6	.02	9.3**
Women	2.3	.01	3.5	.01	22.8	.02	47.7***	0.3	.01	4.1	.01	12.1	.02	32.1***
Cuban Americans														
Men	1.8	.04	9.1	.03	32.2	.04	18.1***	2.8	.03	3.9	.03	15.9	.03	6.4*
Women	1.5	.04	7.6	.03	29.6	.04	15.9***	1.8	.03	3.3	.02	13.1	.03	6.0*
Puerto Ricans														
Men	1.1	.02	4.7	.02	19.3	.04	10.4**	1.2	.02	5.6	.02	3.8	.03	1.3
Women	2.1	.02	4.6	.02	28.1	.03	20.1***	0.7	.01	0.9	.01	5.1	.02	2.7

Note. Logistic regression techniques were used to test age group differences.
*$P < .05$.
**$P < .01$.
***$P < .001$.

Materials and Methods

The HHANES was conducted from 1982 to 1984 by the National Center for Health Statistics. The HHANES employed a complex multistage sampling design aimed at obtaining a representative sample of Mexican Americans in the Southwest (Texas, New Mexico, Colorado, Arizona, and California), Cuban Americans in the Miami area (Dade County), and Puerto Ricans in the New York City area.[8,9] Response rates for the physical examination where audiometric data were collected were 76.1%, 60.8%, and 75%, respectively, for Mexican Americans, Cuban Americans, and Puerto Ricans.[8]

Our analysis is limited to adults aged 20 to 74. Pure tone audiometric data were collected on one half of the subjects participating in the physical examination (Mexican-American n = 1682, Cuban-American n = 441, and Puerto Rican n = 628). All subjects were alternately assigned to audiometric or other medical testing when they scheduled their medical examination. Participants were asked during the physical examination whether they had ever used a hearing aid.

Audiometric testing was conducted in a sound-treated booth using calibrated audiometric equipment.[10,11] Air conduction thresholds were obtained at 500, 1000, 2000, and 4000 Hz, with testing repeated at 1000 Hz. Masked hearing thresholds were obtained using narrow-band noise in the nontest ear when there was an interear difference in air conduction thresholds[11] greater than 40 dB.

Pure tone average thresholds (PTAs) were calculated by averaging thresholds obtained at frequencies of 500 Hz, 1000 Hz, and 2000 Hz.[12] The better of the two threshold responses at 1000 Hz was used in PTA calculation. Masked threshold values were used when indicated. Hearing loss was defined as a PTA of greater than 25 dB in either ear.[12] PTAs were not calculated on subjects with incomplete bilateral audiometric data (n = 28) or when there was probable equipment failure (n = 5).

Analyses were calculated separately for men and women within the three ethnic groups using the general linear model and logistic regression programs from SAS.[13] Age-adjusted percentages were also calculated to make gender-specific comparisons within each of the three ethnic groups.[14,15] Because of the multistage sampling design, analyses were performed with adjustments for sample weights and design effects.[8,9]

Results

Table 1 presents the prevalence rates of hearing loss by age group, ethnicity, and gender. There was a significant increase in the prevalence of hearing loss with increasing age within each ethnic by gender comparison. Prevalence rates ranged from 2% to 7% and 24% to 48% in the youngest and oldest age groups, respectively. Results by ethnicity indicate that, in general, the Puerto Ricans had a lower prevalence of hearing loss.

Table 2 presents the prevalence of mild (26 through 40 dB) and moderate or greater (> 40 dB) hearing loss by age group, gender, and ethnicity. Generally, mild hearing loss was more prevalent than moderate or greater hearing loss. Furthermore, the prevalence of moderate or greater hearing loss exceeded 10% in only the oldest age group. The prevalence of mild hearing loss significantly increased across age groups. The prevalence of moderate or greater hearing loss significantly increased by age group among Mexican Americans and Cuban Americans, but not Puerto Ricans. The prevalence of moderate or greater loss among Puerto Ricans in the 55 to 74 age group was markedly lower than in Mexican Americans and Cuban Americans.

Table 3 presents the age-adjusted prevalence of hearing loss by gender. Hearing loss prevalence was generally higher in men than in women, although these differences were statistically significant only for Mexican Americans when examining overall (> 25 dB) and mild (26 through 40 dB) hearing loss.

Table 4 presents the age-adjusted prevalence of hearing aid use by gender and ethnicity. The prevalence of hearing aid use was less than 10% in each ethnic by gender comparison. Inconsistent nonsignificant gender differences in hearing aid use were noted. The prevalence of hearing aid use among Mexican-American men and women with moderate and greater hearing loss (> 40 dB) was 5.4% (± .05) and 11.1% (± .05), respectively.

TABLE 3—Prevalence of Mild, Moderate, and Overall Hearing Loss by Gender

	Men		Women		
	%	SE	%	SE	χ^2
Mild (26–40 db)					
Mexican Americans	10.5	.01	6.9	.01	7.5[a]
Cuban Americans	15.3	.02	14.1	.02	0.2
Puerto Ricans	6.6	.02	8.4	.01	0.7
Moderate (> 40 db)					
Mexican Americans	5.3	.01	3.9	.01	1.9
Cuban Americans	7.5	.02	6.4	.02	0.3
Puerto Ricans	3.6	.01	1.8	.01	2.3
Overall (> 25 db)					
Mexican Americans	15.8	.01	10.9	.01	10.3[a]
Cuban Americans	22.7	.03	20.4	.02	0.5
Puerto Ricans	10.3	.02	10.2	.01	0.0

Note. Prevalence estimates are age-adjusted. Logistic regression techniques were used to test gender group differences.

[a] $P < .01$.

TABLE 4—Prevalence of Hearing Aid Use by Gender among Hispanics

	Men		Women		
	%	SE	%	SE	χ^2
Overall (> 25 db)[a]					
Mexican Americans	3.7	.02	3.9	.02	0.2
Cuban Americans	1.9	.03	6.7	.03	0.6
Puerto Ricans	9.9	.05	4.9	.05	1.6
Moderate (> 40 db)[b]					
Mexican Americans	5.4	.05	11.1	.05	0.8

Note. Prevalence estimates are age-adjusted.

[a] Denominator includes all participants with > 25 db hearing loss.

[b] Denominator includes all partipants with > 40 db hearing loss. Because of sample size limitations, only results for Mexican Americans are reported. These estimates may be unstable as a result of analytic sample sizes less than 45 (men = 37, women = 38).

Discussion

The present study confirmed that hearing loss—considered to be the second most prevalent impairment in the United States[2]— is also highly prevalent in the Hispanic population. The expected increase in prevalence with advancing age was also observed. A higher prevalence of hearing loss for men, reported elsewhere,[3,7] was confirmed only in Mexican-American men.

The findings by ethnicity suggest that Puerto Ricans have an overall lower prevalence of hearing loss relative to the other two ethnic groups. Possible hypotheses to account for these findings include the following: (1) there are genetic or anthropomorphic characteristics among Puerto Ricans that perform a protective function;[7] and (2) Puerto Ricans in the present study may have experienced less lifetime exposure to damaging levels of noise.[1]

A low prevalence of hearing aid use has been documented in previous research with estimates ranging from 6% to 14%.[3,16] In this report, prevalence rates remained below 12% even when examining Mexican Americans with 40 dB or greater hearing loss. Virtually all individuals with such a loss would benefit from hearing amplification.[17] Factors that may contribute to low hearing aid use include cost and low consumer awareness of the benefits of amplification.[16] Future research should focus on identification of hearing-impaired individuals and the development of culture- and language-specific interventions designed to increase hearing aid use in Hispanics. ☐

Acknowledgments

This work was supported by Grant AG06235 from the National Institute on Aging.

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Interpersonal Violence-Related Injuries in an African-American Community in Philadelphia

Amy R. Wishner, RN, MSN, Donald F. Schwarz, MD, MPH, Jeane Ann Grisso, MD, MSc, John H. Holmes, MS, and Rudolph L. Sutton, MPH

ABSTRACT

Violence has been recognized as a significant health problem. This study describes fatal and nonfatal interpersonal violence-related injury events over 1 year in an indigent African-American community in Philadelphia. Information on injuries was collected from emergency rooms, the Office of the Medical Examiner, and death certificate files. For persons aged 15 through 49 years, violence-related injury rates surpassed any other injury type. The overall violence-related injury rate was 28.7 per thousand population. Interpersonal violence-related injuries were important for the 0- to 4-year age group (9.19 injuries per 1000 population), and continued as a major cause of morbidity through age 59 (12.08 injuries per 1000 population). For more than half of the events, information from the emergency room chart was sufficient only to categorize the incident as a "fight" or that it was intentional; no further classification was possible.(*Am J Public Health.* 1991; 81:1474–1476)

Introduction

Little information is available about the larger problem of violence that underlies homicide in inner-city communities. The current study was carried out by the Philadelphia Injury Prevention Program (PIPP), a cooperative effort of the Philadelphia Department of Public Health, Children's Hospital of Philadelphia, the University of Pennsylvania School of Medicine, the Centers for Disease Control, and a local citizens' advisory board. The study involved fatal and nonfatal interpersonal violence-related injuries (IVRI) in a geographically defined community seen over a 1-year period in emergency rooms, the Office of the Medical Examiner, and death certificate files.

Methods

Study Communities and Case Definition

The neighborhoods in this study comprised 17 census tracts with 68 103 residents in Philadelphia. The population was 97.2% African American, with a median family income of $11 810.

Information on all injuries occurring to residents of the study area was collected by an active emergency room surveillance system, outlined elsewhere.[1] An injury was defined as a case if the event occurred to a resident of the study area and led to emergency room evaluation at a hospital in the injury surveillance network or resulted in death between March 1, 1987, and February 29, 1988.

Trained medical abstractors ascertained from a patient's emergency room chart whether or not an injury occurred and was inflicted by another person in a manner that could be characterized clearly as assaultive or intentional. These injuries were included in the IVRI category. On the basis of the medical record, IVRI were classified as child abuse (cases of abuse up to age 19), rape, stabbing, firearm, or other assault.

Data Collection and Analysis

Data collection, quality control, and analysis have been described previously.[1] Briefly, each injury event was coded according to the International Classification of Diseases, 9th revision, Clinical Modification (ICD-9-CM) External Cause of Injury Codes (E-codes) by an experienced medical coder (see the Appendix for the E-code categories making up each classification). For events with more than one E-code, two of the authors independently assigned a principal code to that event that initiated the sequence leading to physical injury.

Time of entry to an emergency room was classified as morning (6:00 AM to

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Amy R. Wishner and Rudolph L. Sutton are with the Philadelphia Department of Public Health. Donald F. Schwarz and John H. Holmes are with The Children's Hospital of Philadelphia, and Jeane Ann Grisso is with the University of Pennsylvania School of Medicine.

Requests for reprints should be sent to Amy R. Wishner, RN, MSN, Division of Disease Control, Philadelphia Dept. of Public Health, 500 South Broad Street, Philadelphia PA 19146.

This paper was submitted to the journal April 23, 1990, and accepted with revisions April 24, 1991.

11:59 AM), afternoon (12 noon to 5:59 PM), evening (6:00 PM to 11:59 PM), or night (12 midnight to 5:59 AM).

Results

A total of 1958 IVRI were identified. The IVRI rate surpassed the rate for any other injury type for persons aged 15 through 49. For all ages, IVRI had the second highest injury rate per thousand residents (28.7) of any type of injury, less than falls (31.7) and greater than motor vehicle incidents (22.0).

There were 1690 persons who sustained IVRI; 100 individuals had more than one IVRI. Fourteen individuals (7 males and 7 females) had three or more IVRI during the single year of surveillance.[2] (In the *MMWR* article,[2] one injury related to legal intervention and one suicide attempt were included in the definition of repeated violence-related injuries.)

Violence resulted in more deaths than did any other injury type, 44% of all injury deaths, for a death rate of 60.2 per 100 000 population. Among the 64 males and 29 females who died as a result of injury, 32 (50%) and 9 (31%), respectively, died from violence.

The total number of IVRI, by type of violence, can be seen in Table 1.

Injured Persons

IVRI occurred predominantly among those aged 15 through 39 years, who composed 38% of the population and experienced 72.2% of all IVRI. The rate of IVRI among males (40.5 per 1000 males) was approximately twice the rate for females (19.4 per 1000 females). Males constituted 44.3% of this population and experienced 62.4% of the IVRI (Figure 1).

Injury Events

Time of admission to the emergency room was noted in 98.1% (1920) of IVRI. Cases were more likely to enter an emergency room during evening hours (658, or 34.3%) than at night (475, or 24.7%), afternoon (476, or 24.7%), or morning (311, or 16.2%). Approximately one third (620, or 32.1%) of the 1931 IVRI entered an emergency room on a Saturday or Sunday.

The greatest number of IVRI cases entered emergency rooms in the summer months of June (214, or 10.9%), July (187, or 9.6%), and August (180, or 9.2%). The 74 firearm injuries had a different pattern, with the greatest number entering emergency rooms in November and December (11, or 14.9%, and 9, or 12.2%, of firearm injuries, respectively).

Information on alcohol or drug involvement, and where the injury event took place, was rarely recorded. Alcohol involvement was recorded in 7.1% (139) of cases, drug involvement in 1.1% (21), and location in 20.2% (396).

Hospital Admissions. There were a total of 107 hospital admissions, a rate of 5.5 per 100 IVRI. Among males, who had higher hospital admission rates for each category of violence-related injury, firearms ranked first with 28 admissions, 36 for stabbing, and 30 for other assaults. Females had 1 hospital admission for firearms, 2 for stabbing, and 10 for other assaults.

Deaths. Firearm-related injuries had the highest case-fatality rate (CFR) per 100 cases, a sex-specific CFR of 26.5 for males (18 deaths) and 16.7 for females (1 death). Both males and females had a CFR of 3.4 for stabbings, with 8 and 2 deaths, respectively. Other assaults had markedly lower CFRs, 0.8 (6 deaths) for males and 0.9 (4 deaths) for females. Among females, there was 1 death due to a rape and 1 death due to child abuse.

Abuse. There were 63 cases of abuse to children from birth through age 19, a rate of 3.28 per 1000 children.

Rape. Seven rapes were reported, all among females. Six occurred among females aged 15 through 34.

TABLE 1—Total Number of Interpersonal Violence-Related Injuries by Type of Violence

Type of Violence	No. of Events	Rate per 1000 Population
Abuse	63	3.28[a]
Rape	7	0.10
Stabbing	296	4.35
Firearms	74	1.09
Other assaults	1518	22.29
Total	1958	28.71

[a] Rate per 1000 population 0–18 years of age.

Stabbing. Of the 296 stabbings, 80.1% (237) involved male victims. Rates among males peaked in those aged 20 through 24 years, whereas among females stabbing rates peaked among an older age group, those aged 25 through 29 years (Table 2).

Firearms. As with stabbings, firearm injuries increased dramatically beginning in those aged 15 through 19, and among males remained high through ages 40 through 44 (Table 3).

Other Assaults. Assaults not otherwise categorized constituted 77.5% of IVRI, with 1518 events (Table 4). Little information was available about the majority of other assaults; 484 could be classified only as fights, and 578 were not specified as to cause of IVRI.

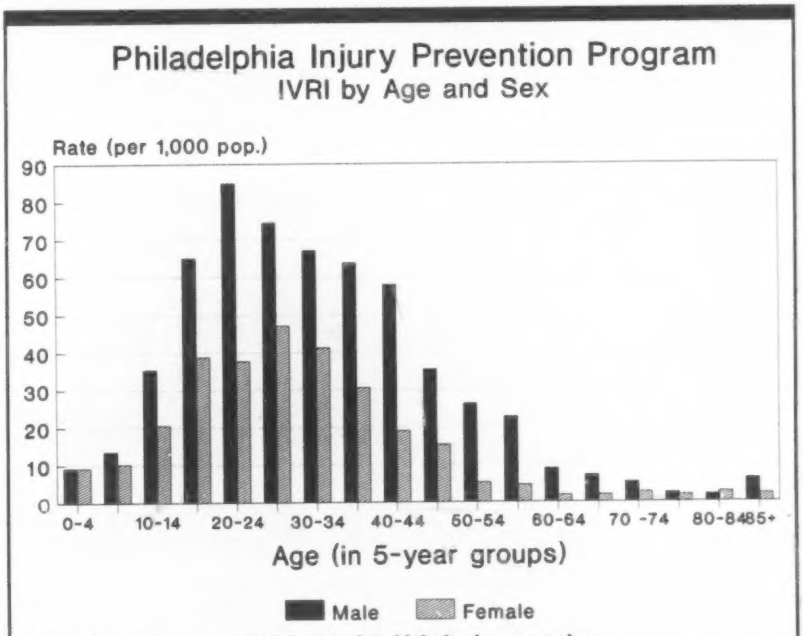

FIGURE 1—Interpersonal violence-related injuries by age and sex.

TABLE 2—Stabbing Rates by Age and Sex (per 1000 population)

Age Group	Males	Females	Total
0–14	0.70	0.44	0.57
15–19	10.33	2.74	6.58
20–24	19.14	2.32	10.52
25–29	17.42	4.70	10.53
30–34	15.40	4.34	9.25
35–39	16.25	4.10	9.34
40–44	12.36	1.12	5.83
45–49	4.31	0.00	4.31
50–54	7.38	1.70	4.03
55+	1.65	0.08	0.67
Total	7.85	1.56	4.35

TABLE 3—Firearm Injury Rates by Age and Sex (per 1000 population)

Age Group	Males	Females	Total
0–14	0.14	0.00	0.07
15–19	4.98	1.56	3.29
20–24	6.61	0.00	3.22
25–29	5.93	0.31	2.89
30–34	2.72	0.36	1.41
35–39	1.81	0.00	0.78
40–44	2.32	0.00	0.97
45–49	0.86	0.00	0.35
50–54	0.00	0.00	0.00
55–59	2.83	0.00	1.18
60–64	0.63	0.00	0.26
65+	0.23	0.00	0.08
Total	2.25	0.16	1.09

TABLE 4—Other Assault Rates by Age and Sex* (per 1000 population)

Age Group	Males	Females	Total
0–4	2.80	1.67	2.25
5–9	11.85	4.43	8.23
10–14	31.52	16.98	24.37
15–19	49.37	32.85	41.20
20–24	59.17	35.06	46.81
25–29	51.15	41.71	46.03
30–34	48.91	35.82	41.63
35–39	45.73	26.44	34.75
40–44	43.28	17.87	28.53
45–49	30.17	15.35	21.37
50–54	18.85	3.41	9.73
55–59	14.84	4.04	8.55
60–64	5.70	1.80	3.42
65–69	5.65	1.84	3.34
70–74	5.22	2.57	3.56
75–79	2.32	1.82	2.00
80–84	1.90	1.71	1.77
85+	6.08	2.16	3.19
Total	29.55	16.46	22.26

*Excludes two other assaults where either age or sex could not be ascertained.

Discussion

These data provide compelling evidence of (1) the magnitude of the problem of violence-related injuries, (2) the need to look beyond homicide and firearm-related data and include stabbings and assaults to appreciate the full spectrum of violence-related injuries, (3) the need to address violence occurring to preschool and elementary school–aged children, and (4) the need for more comprehensive information to be recorded in emergency room charts.

Focusing on fatalities neglects the impact of stabbings and other assaults, more frequent causes of injury to young people.

In PIPP's data, child abuse and other assaults that were seen in the 0 through 4 age group are of particular concern. Observing or being the victim of violence early in life may predispose one to committing or being a victim of violence in adult years.[3–6]

Limitations of the Data

Emergency room records may not always indicate intentional incidents.[7] Thus, the rates reported in this study are conservative estimates. Rape was especially underreported in this study, as the two Philadelphia hospitals with rape treatment programs were not in the surveillance network.

PIPP's study pinpoints needed but inconsistently available data, including alcohol/drug status, victim–perpetrator relationship, time of the injury event, where the injury event took place, characterization of assaults, living arrangements and social support, history of the injury event, previous violence involving the patient either as victim or perpetrator, roles of witnesses and the police, weapon-carrying behavior, and presence of weapons in the household.

Conclusion

The data presented above on fatal and nonfatal IVRI events in a geographically defined African-American community provide dramatic evidence of the staggering extent of violence. We face the urgent task of applying these data in making use of existing violence prevention programs, developing new efforts, evaluating their effectiveness, and improving the data available to help design such projects if they are to have an impact on the important public health problem of violence. ☐

Acknowledgments

This study was supported by a grant from the Centers for Disease Control and the Robert Wood Johnson Foundation Clinical Scholars and General Academic Pediatrics Programs.

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APPENDIX
E-Code Groupings

Abuse

E904.0	Hunger, thirst, exposure, and neglect
E967.0	Child battering and other maltreatment
E968.4	Criminal neglect

Rape

E960.1	Rape

Stabbing

E966	Assault by cutting and piercing instrument

Firearm

E965	Assault by firearms and explosives

Other Assault

E960.0	Unarmed fight or brawl
E961	Assault by corrosive or caustic substance, except poisoning
E962	Assault by poisoning
E963	Assault by hanging and strangulation
E964	Assault by submersion (drowning)
E968	Assault by other and unspecified means
E968.0	Fire
E968.1	Pushing from a high place
E968.2	Striking by blunt or thrown object
E968.3	Hot liquid
E968.8	Other specified means
E968.9	Unspecified means
E969	Late effects of injury purposely inflicted by other person

Note. These groupings exclude suicide or suicide attempts and legal intervention injuries.

Physical Activity and Hypertension in Black Adults: The Pitt County Study

Barbara E. Ainsworth, PhD, MPH, Nora L. Keenan, MPH, David S. Strogatz, PhD, Joanne M. Garrett, PhD, and Sherman A. James, PhD

ABSTRACT

The relation of physical activity to hypertension was examined in 1751 Black adults in Pitt County, NC. More women (65%) than men (44%) were classified as sedentary. Sedentary behavior was not associated with the prevalence of hypertension in men, but was associated with a 31% increase in prevalence for women (sedentary-26.2%; active-20.0%; $P < .01$). The association in women was independent of other risk factors for hypertension. (*Am J Public Health.* 1991;81:1477–1479)

Introduction

Observational and experimental studies suggest that regular physical activity is inversely related to blood pressure.[1–4] However, little is known about physical activity and its relation to blood pressure in Black adults, despite the recognized excess risk of hypertension in Blacks.[5–10] Some surveys of physical activity have detected little difference by race,[11,12] but others have found Blacks to be less active.[10,13–18] In a study of risk factors for heart disease in young adults,[19] duration on a treadmill exercise test was significantly shorter for Blacks than for Whites. None of the reports on physical activity or fitness in biracial samples have race-specific analyses of the association with elevated blood pressure.

The purpose of this article is to describe the level of physical activity and the relation between physical activity and the prevalence of hypertension in a community-based sample of Black adults residing in Pitt County, NC.

Methods

The data for these analyses were collected in the 1988 survey of the Pitt County Study. The purpose of this survey was to provide baseline information for the prospective study of psychosocial and dietary risk factors for elevated blood pressure in Black adults 25 to 50 years of age. Selection of the probability sample and additional information about the study have been published elsewhere.[20] Of 2225 race- and age-eligible persons identified, 1784 (80 percent) were interviewed.

Physical activity was assessed using a modification of the Lipid Research Clinics study questions on physical activity.[21] Participants who reported engaging in exercise or work strenuous enough to sweat for at least 20 minutes, three times per week, were classified as active. Those who reported exercising or laboring less often, less vigorously, or not at all were classified as sedentary.

Three sitting blood pressure measurements were recorded on each respondent's right arm approximately 15 minutes into the interview. Hypertension was defined by an average diastolic blood pressure (average of second and third readings) of 90 mm Hg or greater or by current use of antihypertensive medication. Other predictors of blood pressure measured in this survey and included in the current analyses were age, body mass index (BMI) in kg/m^2, waist to hip ratio (WHR), and alcohol consumption.[20] Sociodemographic descriptors of the study participants were years of education and employment status. Complete data were available on 1751 respondents in the survey.

Weighted analyses were performed to take into account the sampling strategy as well as nonresponse in making population estimates of prevalence. All analyses were gender specific because of expected male-female differences in physical activity and blood pressure. Logistic regression was used to estimate associations between physical activity and hypertension, adjusting for potential confounding by age, BMI, WHR, and alcohol consumption.[22–24]

Results

The distribution and correlates of physical activity are shown in Table 1. Of the 1751 individuals in the analysis, 285 men (44% of 655 men) and 711 women

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Barbara E. Ainsworth is with the Department of Physical Education, Exercise and Sport Science; Nora L. Keenan, David S. Strogatz, and Joanne M. Garrett are with the Department of Epidemiology, School of Public Health; and Joanne M. Garrett also is with the Health Services Research Center—all at the University of North Carolina, Chapel Hill. David S. Strogatz is also with the School of Public Health, State University of New York at Albany and New York State Department of Health, Albany, NY. Sherman A. James is with the Institute for Social Research, and the Department of Epidemiology, School of Public Health, University of Michigan, Ann Arbor.

Requests for reprints should be sent to Barbara E. Ainsworth, PhD, MPH, Department of Physical Education, Exercise and Sport Science, 025 Fetzer Gym, CB 8700, University of North Carolina, Chapel Hill, NC 27599-8700.

This paper was submitted to the journal January 29, 1991, and accepted with revisions June 19, 1991.

TABLE 1—Distribution and Characteristics of 25- to 50-Year-Old Black Men and Women, by Sex and Physical Activity Status: Pitt County, NC, 1988 (N = 1751)

Sex and Characteristics	Sedentary	Active	Difference[a]	95% CI
Men (n = 655)	285	370		
Unemployed	12.4%	4.3%	+8.1%	+4.2–+12.0
<12 years education	29.3%	31.8%	−2.5%	−10.3–+5.3
Mean				
Age	35.6	34.7	+0.9	−0.4–+2.2
BMI	26.4	26.3	+0.1	−0.8–+1.0
WHR	0.899	0.883	+0.016	+0.006–+0.026
Drinks/week	4.2	4.1	+0.1	−1.4–+1.6
Women (n = 1096)	711	385		
Unemployed	22.4%	11.7%	+10.7%	+5.2–+16.2
<12 years education	26.9%	34.0%	−7.1%	−12.0–−2.2
Mean				
Age	35.3	35.7	−0.4	−1.5–+0.7
BMI	29.6	29.0	+0.6	−0.4–+1.6
WHR	0.852	0.851	+0.001	−0.013–+0.015
Drinks/week	1.2	1.3	−0.1	−0.7–+0.5

Note. CI = confidence interval; BMI = body mass index; WHR = waist to hip ratio.
[a]Difference = Sedentary − Active.

TABLE 2—Weighted Multivariable Logistic Regression Coefficients and Standard Errors for the Prediction of Prevalence of Hypertension: 25- to 50-Year-Old Black Women and Men, Pitt County, NC, 1988

	Men		Women	
Variables	β	SE	β	SE
Sedentary behavior	−0.203	0.188	0.466	0.171
Age	0.075	0.014	0.098	0.012
BMI	0.119	0.024	0.081	0.012
WHR	2.740	1.539	5.028	1.124
Alcohol consumption (drinks/week)				
0 (referent)	—	—	—	—
<1	0.427	0.292	0.243	0.227
1–6	0.006	0.231	0.449	0.229
≥7	0.793	0.256	0.888	0.361

Note. β = logistic regression coefficient; SE = standard error; BMI = body mass index; WHR = waist to hip ratio.

(65% of 1096 women) were classified as sedentary. Sedentary adults of both sexes were more likely to be unemployed. A weaker association was observed between sedentary behavior and being a high school graduate. Age, BMI, and consumption of alcohol were unrelated to physical activity. Among men, WHRs were slightly greater among those with sedentary life-styles.

Prevalence of hypertension was higher in the sedentary than in the active women (26.2% vs 20.0%, $P < .01$), but there was no significant difference between sedentary and active men (30.6% vs 31.7%, respectively). These findings persisted (Table 2) after adjustment for age, BMI, WHR, and alcohol consumption. The adjusted odds ratios (95% confidence intervals) were 1.59 (1.14–2.34) and 0.82 (0.57–1.18) for women and men, respectively. The results did not change when education and unemployment were added to the model.

Discussion

In this sample of Black adults, the majority of the women and a sizable minority of the men were classified as sedentary. The difference by sex is consistent with reports from other cross-sectional surveys.[13,15,18,25,26] The overall level (57%) of sedentary behavior in this population of young to early middle-aged adults was higher than expected, nearly equaling the median level (59 percent) observed in the 1987 Behavioral Risk Factor Surveillance System,[27] a survey of 32 states that included older, presumably less active adults.

The prevalence of hypertension was significantly higher in self-reported sedentary adults only among women. A similar difference was reported in cross-sectional findings on blood pressure from the Stanford Community Health Survey,[25] but the results from studies of men and women generally are not consistent with sex-specific effects.[19,28-32] Two possible explanations for the difference in findings for men and women were examined in these data. The association between exercise and hypertension in men could have been biased toward the null if previously sedentary men were especially likely to begin exercising after being diagnosed as hypertensive. However, among hypertensive adults no association was observed between activity and awareness of hypertension, making this form of bias an unlikely explanation for our results. The temporal relation between physical activity and hypertension cannot be determined with complete certainty in these data. A second possibility is that the results for men and women reflect interaction between obesity and activity, with activity being more protective in the presence of obesity, which was far more common in women. This pattern of interaction was reported for Harvard alumni,[33] but was not present in sex-specific stratified analyses of these data. Nondifferential misclassification of activity cannot be evaluated with available data and may have been greater for men than for women.

Most recent surveys have found a direct cross-sectional association between indicators of socioeconomic status and physical activity.[18,34] The weak inverse association between education and physical activity in this population may be a chance finding or may be due in part to job-related activity required of those with less education in Pitt County. The association between unemployment and sedentary behavior is consistent with the notion of work as a source of physical demands or a source of financial means for engaging in leisure activity. However, the overall level of sedentary behavior in this community suggests that increases in physical activity, even among the employed, are needed. The high prevalence of reported sedentary behavior also suggests that an understanding of community-level influences on exercise and activity may be necessary for successful intervention.[35,36] □

Acknowledgments

This research is supported by grant HL33211 from the National Heart, Lung, and Blood Institute. The authors thank Janet Croft, Ann Marie Gerber, Robert Kleckner, and Brian Rice for their contributions to this research and to Cindy Atkins for typing the manuscript.

This paper was presented at the 118th Annual Meeting of the American Public Health Association, October 3, 1990, New York City.

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Black-White Differences in Infant Mortality in 38 Standard Metropolitan Statistical Areas

ABSTRACT

The Black-White difference in infant mortality rates for 1982 through 1986 in 38 large US standard metropolitan statistical areas (SMSAs) varied by a factor of almost seven. In multiple regression analyses the most important predictor of the Black-White difference in the 38 SMSAs was an index of Black-White residential dissimilarity (or "segregation index"), independent of Black-White differences in median family income and poverty prevalence. Certain SMSAs in California had relatively low segregation indexes and small Black-White differences in infant mortality, despite considerable Black-White differences in poverty prevalence. The explanations for the apparent effect of residential segregation should be explored. (*Am J Public Health.* 1991;81:1480-1482)

Anthony P. Polednak, PhD

Introduction

The US Black-White difference in the infant mortality rate has persisted. Lower rates have been reported for Blacks in the West than in the other three US census regions.[1] This report examines the Black-White difference in infant mortality rate in 38 large standard metropolitan statistical areas (SMSAs), including several in California, in relation to socioeconomic status indicators and an index of residential segregation. Residential segregation may be associated with such factors as the availability and quality of prenatal and postnatal medical care, which may influence infant mortality independent of differences in socioeconomic status.

Methods

For each of 38 SMSAs with a total population of more than 1 million in 1980, 1980 census reports were used to obtain Black-White differences in poverty prevalence, median family income, and the percentage of families with a "female householder, no husband present."[2] One segregation index (an index of residential dissimilarity between Blacks and Whites in the SMSAs based on 1980 census data) measures the unevenness of residential distribution of a specific minority population across census tracts within an urban area. The formula is as follows:

$$0.5 \times \sum_{i=1}^{n} \mid (x_i/X) - (Y_i/Y) \mid ,$$

where x and y are the numbers of each race with a tract (i) and X and Y are the total populations of each race in the SMSA.[3] Numbers of deaths at less than 1 year of age for Blacks and Whites in each SMSA for 1982 through 1986 were used,[4] along with total live births in these years,[5] to obtain an average annual infant death rate (per 1000 live births) for Blacks and Whites in each of the 38 SMSAs.

Results

Infant mortality rates were higher for Blacks than for Whites in all of the 38 SMSAs. The mean Black-White difference was 8.63 per 1000 live births (SD = 3.27), ranging from 2.14 (95% confidence limits [CL] = −0.90 and 5.06) for Anaheim, Calif, to 14.63 (95% CL = 12.31 and 16.95) for Pittsburgh, Pa (see Appendix for detailed data).

The segregation index and the Black-White difference in poverty prevalence for the 38 SMSAs were significantly correlated ($r = .573$, $P < .001$), but this correlation does not indicate strong collinearity, and both variables were included in multiple regression analysis (Table 1). The segregation index was the only statistically significant independent predictor of the Black-White difference in infant mortality rate among the SMSAs, and the only independent variable selected in stepwise regression (not shown). Only for Blacks was the segregation index a statistically significant independent predictor of infant mortality rate in multiple regression analysis (data not shown).

Chicago had the highest segregation index (i.e., .878) and, a Black-White difference in poverty prevalence (i.e., 28.8%) that was higher than the average (i.e., 18.3%), the Black-White difference in infant mortality was also large (i.e., 13.38 per 1000; 95% CL = 12.39 and 14.17). The seven California SMSAs differed considerably in segregation index. Black-White differences in poverty prevalence ranged from 7.8% for San Jose to 16.6% for Sacramento and 17.4% for Riverside; all were

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The author is with the Department of Epidemiology and Public Health, Yale School of Medicine. Part of this work was done while he was with the Department of Preventive Medicine, School of Medicine, State University of New York at Stony Brook, NY.

Requests for reprints should be sent to Anthony P. Polednak, Department of Epidemiology and Public Health, Yale School of Medicine, 60 College Street, PO Box 3333, New Haven CT 06510.

This paper was submitted to the journal October 5, 1990, and accepted with revisions May 28, 1991.

below the average for all 38 SMSAs. Anaheim had the lowest segregation index and the smallest Black–White difference in infant mortality, whereas Los Angeles had the highest segregation index and the largest Black–White difference in infant mortality (Figure 1), despite the similarity in the Black–White difference in poverty prevalence in the two areas (i.e., 9.9% and 8.7%, respectively).

Discussion

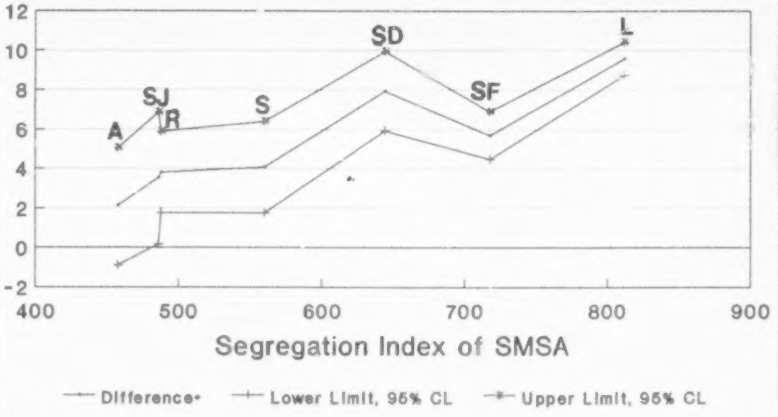
The almost sevenfold variation among the 38 SMSAs in the Black–White difference in infant mortality rate and the association with the segregation index (apparently independent of variation in the Black–White difference in poverty prevalence) require explanation. Socioeconomic status may not have been adequately controlled for in the analysis of the effect of the segregation index on the Black–White difference in infant mortality; areas with high segregation indexes may include subareas (in inner cities) of extreme poverty and high cost of living. Potential inaccuracies in infant death rates by SMSA due to differential rates (among the SMSAs) of migration of mothers (with their infants) after birth should be examined by separate analysis of neonatal and postneonatal death rates (both of which are higher in US Blacks than Whites). Neonatal death rates would be little affected by migration, because migration during this period is unlikely. Linked live birth–infant death files also would be useful.

Despite considerable Black–White differences in poverty prevalence, two SMSAs in California (i.e., Riverside and Sacramento) with low indexes of segregation had small Black–White differences in infant mortality (Figure 1). The explanation for low Black infant mortality rates in Anaheim, with both a small Black–White difference in poverty prevalence and a low segregation index, also should be explored. Birth weight distributions, birth weight–specific death rates, and both neonatal and postneonatal death rates should be examined. Binkin et al.[6] suggested that the smaller difference in Black–White neonatal death rates within normal weight births in California than in Georgia could reflect better quality and availability of prenatal, intrapartum, and postnatal care in California Blacks. The need for studies of the quality vs the quantity of prenatal care in Blacks has been recognized,[7] and the same holds for postnatal care. SMSAs with high levels of segregation are mainly large, older industrial cities with Black ghettos.[3] Among the many potential explanations for higher infant mortality rates in these segregated areas, availability (or accessibility) and use of diagnostic/treatment procedures, level of training and attitudes or recommendations of providers, and patient (i.e., maternal) decision making should be examined, as suggested for explaining racial inequalities in use of diagnostic/treatment procedures for cardiovascular disease.[8] ☐

TABLE 1—Prediction of the Black–White Difference in Infant Mortality Rates (1982–1986) in 38 SMSAs

	Regression Coefficient	*t*	*P* Values
Female householder[a]	−.0672	−.263	.794
Poverty prevalence[b]	.2503	1.524	.137
Median family income[c]	−.1918	−1.054	.299
Segregation index[d]	.7469	4.033	<.001

Note. SMSA = standard metropolitan statistical area.
[a]Black–White difference in proportion of families with "female householder, no husband present" based on 1980 US census reports.
[b]Black–White difference in proportion of persons below the poverty level in 1979, based on 1980 US census reports.
[c]Black–White difference in median family income, based on 1980 US census reports.
[d]Black–White residential segregation index for each SMSA, as reported by Massey and Denton.[3]



•Black-white difference in infant mortality rate

FIGURE 1—Infant mortality difference in California standard metropolitan statistical areas (SMSAs). (A = Anaheim; L = Los Angeles; R = Riverside; S = Sacramento; SD = San Diego; SF = San Francisco; SJ = San Jose.)

Acknowledgments

The author thanks two anonymous reviewers for critical comments on the potential effects of differential migration on infant death rates by race/SMSA and on the need for discussing potential mechanisms for segregation effects.

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APPENDIX—Segregation Index, Prevalence of Poverty, and Infant Mortality Rates for 38 SMSAs

		Prevalence of Poverty (%)			Infant Mortality Rate per 1000 Live Births (1982–1986)			
SMSA	Segregation Index[a]	Black	White	Difference	Black Rate	White Rate	Black–White Difference	95% CI[b]
Anaheim, Calif	.458	15.84	5.99	9.85	10.59	8.45	2.14	−0.90, 5.06
Atlanta, Ga	.785	26.02	10.96	15.06	17.36	9.32	8.04	6.85, 9.23
Baltimore, Md	.747	27.02	6.99	20.03	18.50	9.19	9.31	8.05, 10.57
Boston, Mass	.776	25.86	8.81	17.05	18.45	8.45	10.00	8.21, 11.79
Buffalo, NY	.794	33.09	7.20	25.89	17.70	8.70	9.00	6.52, 11.48
Chicago, Ill	.878	34.46	5.71	28.75	23.03	9.65	13.38	12.59, 14.17
Cincinnati, Ohio	.723	32.78	6.79	25.99	18.40	9.53	8.87	6.89, 10.85
Cleveland, Ohio	.875	28.65	6.33	22.32	19.25	8.83	10.41	8.85, 11.97
Columbus, Ohio	.714	23.52	8.46	15.06	15.74	10.52	5.21	3.08, 7.34
Dallas, Tex	.771	25.13	6.38	18.75	17.39	9.35	8.05	6.90, 9.20
Denver, Colo	.684	18.36	6.90	11.46	15.42	9.39	6.03	3.54, 8.52
Detroit, Mich	.867	28.97	6.93	22.04	23.38	9.42	14.41	13.28, 15.56
Ft Lauderdale, Fla	.816	29.32	6.20	23.12	21.05	7.99	13.06	10.88, 15.24
Houston, Tex	.695	22.44	6.20	16.24	15.34	9.24	6.09	5.07, 7.11
Indianapolis, Ind	.762	28.50	6.80	21.70	23.81	10.52	13.29	10.89, 15.69
Kansas City, Kan-Mo	.789	25.19	6.27	18.92	17.13	9.68	7.45	5.57, 9.33
Los Angeles, Calif	.811	18.06	9.36	8.70	18.71	9.14	9.57	8.72, 10.42
Miami, Fla	.778	30.64	10.29	20.35	17.32	7.75	9.57	8.27, 10.87
Milwaukee, Wis	.839	35.34	5.47	29.87	17.97	8.41	9.57	7.72, 11.42
Minneapolis, Minn	.683	26.50	5.33	21.17	16.80	8.73	8.07	5.36, 10.78
Nassau-Suffolk, NY	.755	16.98	4.65	12.33	22.19	7.88	14.31	12.01, 16.61
New Orleans, La	.683	37.67	9.30	28.37	18.61	9.37	9.24	7.80, 10.68
New York, NY	.820	28.62	10.79	17.83	16.22	10.88	5.34	4.73, 5.95
Newark, NJ	.816	29.02	6.45	22.57	19.60	8.21	11.39	9.92, 12.86
Philadelphia, Pa	.788	30.19	7.06	23.13	20.60	9.44	11.16	10.16, 12.16
Phoenix, Ariz	.594	18.39	8.26	10.13	14.51	9.07	5.44	2.81, 8.07
Pittsburgh, Pa	.727	27.39	6.34	21.05	24.07	9.43	14.63	12.31, 16.95
Portland, Ore	.685	22.86	7.76	15.10	14.19	9.54	4.66	1.15, 8.17
Riverside, Calif	.488	26.69	9.34	17.35	14.45	10.63	3.82	1.76, 5.88
Sacramento, Calif	.559	26.14	9.59	16.55	12.85	8.78	4.07	1.75, 6.39
Saint Louis, Mo	.813	29.69	6.23	23.46	19.66	8.58	11.08	9.74, 12.42
San Antonio, Tex	.636	29.08	16.27	12.81	16.45	10.54	5.90	3.05, 8.75
San Diego, Calif	.643	18.10	9.44	8.66	16.75	8.82	7.93	5.92, 9.94
San Francisco, Calif	.717	17.63	7.13	10.50	14.41	8.73	5.68	4.47, 6.89
San Jose, Calif	.487	14.45	6.62	7.83	12.21	8.69	3.53	0.19, 6.87
Seattle, Wash	.682	20.57	6.66	13.91	16.68	9.40	7.28	4.34, 10.22
Tampa, Fla	.726	35.06	9.37	25.69	20.12	10.45	9.66	7.58, 11.74
Washington, DC	.701	17.11	4.52	12.59	20.08	8.91	11.17	10.12, 12.22

Note. SMSA = standard metropolitan statistical area.
[a] From Massey and Denton.[3]
[b] Confidence interval (CI) on the difference in rates.

Comparison of Smoking Habits of Blacks and Whites in a Case-Control Study

Geoffrey C. Kabat, PhD, Alfredo Morabia, MD, PhD, and Ernst L. Wynder, MD

ABSTRACT

Information from Blacks and Whites interviewed in a case-control study of tobacco-related diseases was analyzed to identify explanatory factors for racial differences in smoking habits. Blacks were three times more likely to be light vs heavy smokers. This association did not differ according to such variables as cigarette preference, degree of inhalation, or quitting. The association of race and light smoking was present in both current and ex-smokers. Sociodemographic or smoking-related characteristics do not appear to explain racial differences in smoking habits. Future studies should focus on cultural factors influencing smoking behavior. (*Am J Public Health.* 1991; 81:1483–1486)

Introduction

Incidence rates of tobacco-related cancers, such as cancers of the lung, esophagus, larynx, and oral cavity, are higher in Black than in White Americans.[1] In addition, Blacks are more often current smokers but, on average, smoke fewer cigarettes per day than Whites.[2–8] Blacks also tend to smoke higher tar, higher nicotine, and more mentholated cigarettes than Whites.[5,9,10]

Understanding the causes of racial differences in smoking habits has major public health implications since it determines the best strategy for smoking cessation. Racial differences are apparently not attributable to socioeconomic factors[2] but, to our knowledge, the role of smoking-related factors such as tar level, mentholation, inhalation, degree of habituation, butt length, and quitting has not been examined heretofore.

Methods

Data presented here derive from a previously described hospital-based case–control study of tobacco-related cancers.[11] Briefly, case patients with cancers of the lung, larynx, oral cavity, esophagus, bladder, pancreas, and kidney and first occurrences of nonfatal myocardial infarction were interviewed in 28 collaborating hospitals in eight US cities. At least one control patient with a condition not known to be related to smoking or alcohol was interviewed within 2 months following the case patient's interview. Control patients were matched to case patients on sex, age (+/− 5 years), race, and hospital.

A total of 23 011 case and control patients, of whom 9% were Black, were interviewed between 1980 and 1990. After removal of exclusively pipe and/or cigar smokers (n = 903), 5301 never smokers (8% Black), 9252 current smokers (11% Black), and 7555 former smokers (6% Black) remained for analysis.

The distribution of diagnoses was similar between Black and White case patients. The most common case diagnoses in men were lung cancer (37% of White and Black case patients), myocardial infarction (21% of Whites, 18% of Blacks), oral cancer (13%, 17%), and larynx cancer (7%, 8%). Among control patients, Blacks had a smaller proportion of cancers than Whites, because Blacks are underrepresented in collaborating hospitals with almost exclusively cancer patients.

The study questionnaire contained detailed items on smoking history, including age at initiation; tobacco products (cigarettes, cigars, pipes, chewing tobacco, snuff); up to four cigarette brands, specifying whether filter or mentholated; years of smoking each brand; and, for ex-smokers, the number of years since quitting. Information was also available on inhalation (not at all, slightly, moderately, deeply), how much of the cigarette the subject usually smoked, and how long after waking the subject generally smoked his or her first cigarette of the day (0 through 15 minutes, 16 through 30 minutes, 31 through 59 minutes, 60+ minutes).

A smoker was defined as one who had smoked at least one cigarette per day for at least 1 year. A smoker who had smoked within a year of diagnosis was a current smoker, otherwise an ex-smoker. A "heavy" smoker was defined as a smoker of more than a pack per day (≥ 21 cigarettes per day). Tar level was categorized into low, moderate, and high tar (< 10, 10 through 14.9, ≥ 15 mg/cigarette). Mentholation was categorized into nonmenthol only (those who had never smoked a mentholated brand for at least 1 year), menthol

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At the time of this study, all authors were with the American Health Foundation, New York. Alfredo Morabia is now with the Clinical Epidemiology Unit of the University Cantonal Hospital, Geneva, Switzerland.

Requests for reprints should be sent to Geoffrey C. Kabat, PhD, Division of Epidemiology, American Health Foundation, 320 East 43rd Street, New York, NY 10017.

This paper was submitted to the journal April 5, 1991, and accepted with revisions July 9, 1991.

TABLE 1—Distribution of Smoking Status by Case-Control Status, Sex, and Race

	Men (%)		Women (%)	
Smoking Status	Whites[a]	Blacks[b]	Whites[c]	Blacks[d]
Case patients				
Never smoked	9.9	7.8	23.4	23.5
Current smoker	52.7	67.9	55.2	59.4
Ex-smoker	37.4	24.3	21.4	17.1
Control patients				
Never smoked	25.3	22.2	52.2	56.2
Current smoker	29.7	49.1	23.8	29.5
Ex-smoker	45.0	28.7	24.0	14.3

[a]n = 8998 case patients and 6755 control patients.
[b]n = 666 case patients and 562 control patients.
[c]n = 3140 case patients and 3300 control patients.
[d]n = 345 case patients and 342 control patients.

TABLE 2—Comparison of Smoking Characteristics of White and Black Control Patients by Sex (current smokers only)

	Men (%)		Women (%)	
	Whites (n = 2007)	Blacks (n = 278)	Whites (n = 784)	Blacks (n = 101)
Amount Smoked				
1–5 cpd	4.7	10.1	8.4	14.8
6–10 cpd	9.3	26.5	14.8	34.6
11–15 cpd	4.7	17.0	7.0	10.9
16–20 cpd	32.6	30.8	38.8	24.8
21+ cpd	48.7	15.6	31.0	14.9
Tar level				
< 10 mg	18.9	8.4	28.1	13.6
10–14.9 mg	23.0	10.9	33.9	20.4
≥ 15 mg	58.1	81.7	38.0	66.0
Menthol				
Nonmenthol only	79.2	64.0	77.3	48.5
Menthol 1–14 y	12.6	17.0	12.8	21.8
Menthol ≥ 15 y	8.2	19.0	9.9	29.7
Duration of smoking				
< 30 y	27.1	33.7	34.3	43.0
≥ 30 y	72.9	66.3	65.7	57.0
Proportion of cigarette smoked				
All	27.6	28.4	31.1	38.4
⅞–¾	46.5	42.2	34.4	29.5
½	22.6	26.9	30.0	27.7
¼–½	3.3	2.5	4.5	4.4
Inhalation				
Not at all/slightly/ moderately	73.5	78.6	84.6	94.1
Deeply	26.5	21.4	15.4	5.9
Timing of first cigarette				
0–15 minutes	48.8	45.5	39.0	40.7
16–30 minutes	18.7	20.1	21.5	21.2
31–59 minutes	15.9	14.3	16.4	12.4
60+ minutes	16.6	20.1	23.1	25.7

Note. Control patients were hospital patients with non–tobacco-related conditions. cpd = cigarettes per day.

1 through 14 years, or menthol 15 years or more based on lifetime smoking history.

We present Black–White comparisons in case and control patients to demonstrate that the pattern of differences is not an artifact of the study design. We also determined whether differences in smoking between the two races were present in both current and ex-smokers. We computed the odds ratio (OR) of being a light vs a heavy smoker related to race (Black vs White) and used logistic regression to adjust for confounding variables and to assess multiplicative interaction.

Results

The breakdown of the total study population by case–control status, sex, and race is displayed in Table 1. As in national surveys, the prevalence of current smoking was higher in Blacks. Among current smokers in the control group, Blacks smoked fewer cigarettes per day, tended to smoke higher tar and mentholated cigarettes, and tended to have smoked for fewer years compared with Whites (Table 2). Black men and women were much more alike in terms of amount smoked than were White men and women (Table 2).

Among control patients (not shown in Table 2), Black men were 5.0 (95% confidence interval [CI] = 3.3–7.1) times more likely, and Black women were 2.5 (95% CI = 1.4–6.7) times more likely, to be light smokers. For case patients, the corresponding ORs and 95% CIs were 3.0 (2.6–4.2) for men and 3.0 (2.0–4.4) for women. These associations persisted across strata of age, education, marital status, Quetelet's Index, menthol use, tar level, inhalation, duration of smoking, butt length, and timing of first cigarette.

As shown in Table 3, among male control patients the association of race with light smoking was greater in current smokers (OR = 5.0; 95% CI = 3.4–7.5) than in ex-smokers (OR = 2.5; 95% CI = 1.6–3.7). The interaction was statistically significant (*P* = .01) and was of similar magnitude within strata of age, Quetelet Index, tar level, inhalation, and alcohol. Differences between current and ex-smokers were even stronger among more educated subjects and among smokers of mentholated cigarettes. In these two subgroups, among current smokers Blacks were about 7 times more likely than Whites to be light smokers, but among ex-smokers Blacks were only about 1.5 times more likely to be light smokers. No interaction among smoking status, level of smoking, and race was seen in female control patients (not shown in Table 3).

Discussion

In the present study, more Blacks smoked; however, among smokers, Blacks smoked fewer cigarettes per day than Whites. This pattern was constant across subgroups defined by sex, case–control status, sociodemographic factors, life-style factors, and smoking behaviors, and is consistent with US survey data.[2,4,6]

Novotny et al. reported that Blacks were less likely than Whites to be heavy smokers, after adjustment for employment status, blue/white collar status, education, and poverty level.[2] Our results are consistent with the latter findings but, in addition, suggest that the larger proportion of light smokers among Black current smokers is not explained by differences in smoking-related factors.

The agreement of our results with National Health Interview Survey data makes selection or recall bias of hospital patients an unlikely explanation.[4,12] However, purchasing power could account for the racial difference in level of smoking. Although, in the present study, Black–White differences persisted within strata of educational level, education may be a poor indicator of socioeconomic status when comparing Blacks and Whites.

Another explanation is that since the proportion of quitters is higher among Whites and since light smokers are more likely to quit,[13] this could lead to an overrepresentation of heavy smokers among White current smokers. Indeed, there is such a phenomenon, but it does not completely explain the observed association of race and smoking habits since Black–White differences persist, even though attenuated, among ex-smokers (Table 3). It is also possible that Blacks are less susceptible to becoming habituated to smoking since they inhale less (Table 2) and have higher mean cotinine levels compared with Whites for a similar level of smoking.[14,15]

In conclusion, smoking-related factors do not seem to explain racial differences in smoking habits. Thus, effective preventive strategies aiming at smoking cessation may require a better understanding of cultural factors that affect smoking, such as the timing of the most "important" cigarettes, whether smoking is engaged in mainly in the company of others, reason for smoking, reason for smoking mentholated cigarettes (where applicable), and degree of satisfaction from smoking. □

TABLE 3—Adjusted Odds Ratios of Being a Light Smoker for a Black Compared with a White, by Smoking Status, in Male Control Patients, 1980–1990

	Current Smokers	Ex-Smokers	*P* value[a]
Overall	5.0	2.5	.01
Age			
< 60 yrs	4.7	2.6	.11
≥ 60 yrs	5.3	2.4	.12
Education			
< 13 yrs	4.3	3.0	.26
≥ 13 yrs	7.1	1.6	.01
QI			
< 26	4.5	2.4	.06
≥ 26	5.8	2.7	.13
Tar level			
< 15 mg	4.5	2.5	.05
≥ 15 mg	5.5	2.4	.21
Menthol			
Current menthol	7.7	1.3	.01
Current nonmenthol	4.2	2.7	.15
Inhalation			
Deeply	4.0	2.6	.45
Not at all/light/medium	5.6	2.3	.006
Alcohol			
None/occasionally/light	4.1	2.4	.20
Heavy	5.3	2.6	.07

Note. QI = Quetelet's Index. Odds ratios were adjusted for all variables other than the stratifying variable. Variables were categorized as follows: amount smoked (1–20, ≥ 21 cigarettes per day), mentholation (nonmenthol only, menthol 1–14 years, menthol ≥ 15 years), inhalation (deeply, else), timing of first cigarette of the day (≥ 60, < 59 minutes after waking), age (< 60, ≥ 60), education (< 13, ≥ 13 years), alcohol (never/occasionally, ≥ 1 oz whiskey equivalents per day), QI (≥ 26, < 26), smoking duration (< 30/≥ 30 years), portion of cigarette smoked (all, ≤ ¾), and type of hospital. The model also included a variable indicating whether the subject was a current or ex-smoker and a product term between this variable and amount smoked (< 20, ≥ 20 cigarettes per day). Control patients were hospital patients with non–tobacco-related conditions.

[a] Test for homogeneity of the odds ratios between current and ex-smokers.

Acknowledgments

Portions of the material contained in this paper were presented at the 12th scientific meeting of the International Epidemiological Association, Los Angeles, August 9, 1990.

This research was supported by National Cancer Institute Contract N01-CP-05684 and Grant CA-32617.

We would like to acknowledge the valuable contributions of the following cooperating institutions and individuals: Memorial Hospital (New York, NY), Dr. David Schottenfeld and Dr. Nael Martini; Manhattan Veterans Hospital (New York, NY), Dr. Norton Spritz; Long Island Jewish Hillside Medical Center (New York, NY), Dr. Arthur Sawitsky; University of Alabama Hospital (Birmingham, Ala), Dr. William Bridgers; Birmingham Veterans Hospital (Birmingham, Ala), Dr. Herman L. Lehman; Loyola University Hospital (Chicago, Ill), Dr. Walter S. Wood; Hines Veterans Hospital (Chicago, Ill), Dr. John Sharp; Henry Ford Hospital (Detroit, Mich), Robert M. O'Bryan, MD, and Christine Johnson, PhD; Hospital of the University of Pennsylvania (Philadelphia, Pa), Dr. Paul Stolley; Jefferson Medical College and Thomas Jefferson University Hospital (Philadelphia, Pa), Dr. J. E. Colberg; Allegheny General Hospital (Pittsburgh, Pa), Dr. Stanley A. Briller; University of Pittsburgh Eye and Ear Hospital (Pittsburgh, Pa), Dr. Lewis H. Kuller; Pittsburgh Veterans Hospital (Pittsburgh, Pa), Dr. Eugene N. Meyers; Moffitt Hospital and University of California at San Francisco and County Hospital (San Francisco, Calif), Dr. Nicholas Petrakis; St. Luke's Hospital (San Francisco, Calif), Dr. Richard A. Bohannan; Georgia Baptist (Atlanta, Ga), Dr. A. H. Letton; and Emory University Clinic (Atlanta, Ga), Dr. K. Mansour. We would also like to thank Dr. Ken Resnicow for thoughtful comments on the manuscript, Pao Huan Chung for programming assistance, and Laurie Marmelstein for preparation of the manuscript.

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A 5-Year Follow-up Study of 117 Battered Women

Bo Bergman, MD, DrMed Sci, and Bo Brismar, MD, DrMed Sci

ABSTRACT

This paper reports register data concerning somatic and psychiatric hospital care on 117 battered women who were identified in a surgical emergency department and offered a treatment program. Data were collected during a period of 10 years before to 5 years after the battering in question. It was concluded that the battered woman seeks hospital care much more than the average woman of the same age. It is, however, not only traumatic injuries that bring her to the hospital, but also medical, gynecological, psychiatric, and unspecified disorders and suicide attempts. In this study it was hypothesized that this overuse of hospital care reflects the situation at home characterized by ongoing battering and other psychosocial problems. During the 5 years following the battering, the women did not show any signs of reducing their use of hospital care. It is alarming that this high use of medical care continues over years, and doctors should consider battering as one possible explanation for this phenomenon.(*Am J Public Health.* 1991;81: 1486–1489)

Introduction

During the last few decades, with more and more attention being given to the problem of wife battering, various intervention programs aimed at stopping the battering or otherwise helping or supporting battered wives have been tried.[1–3] However, conclusive information concerning the effect of different intervention programs is sorely lacking.

In this study the use of medical care will be used as an indicator of the general life strain, including battering, under which battered women live. The aim of the study was to investigate the use of somatic and psychiatric hospital care by battered women and to use these data as a model to evaluate the effectiveness of a wife battering intervention program in a medical setting.

Subjects and Method

From November 1983 to June 1984, there was an emergency room project at the Huddinge Hospital, Huddinge, Sweden, that attempted to identify and support battered women seeking hospital care. All women in the emergency room who reported injuries resulting from battering were offered admission to the hospital even if the injuries did not warrant it per se. The women were given supportive counseling by a social worker and, if necessary, they talked to a psychiatrist. They were also offered continued supportive counseling by the social worker after returning home, and, on these occasions, they were provided an opportunity to contact social services, lawyers, and other sources of help.

During the 8-month period in question, a total of 117 battered women sought help in the emergency room. Fifty-eight agreed to enter the treatment program, and 59 went home after routine care. The batterer was the woman's husband, cohabitant, or steady partner (65%); her ex-husband or ex-cohabitant (19%); or a close male relative (16%). The mean age at the time of the battering was 33 ± 2 years (range: 16–75). The battered women were compared with a control group that was selected from the population register and matched with the group of 117 battered women for age and geographic area in Stockholm.

The battered women and the control subjects were studied regarding consumption of somatic and psychiatric hospital care during the period from 10 years before to 5 years after the battering. These data were obtained from the Stockholm County Council's computer files, supple-

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Bo Bergman is with the Department of Psychiatry and Bo Brismar is with the Department of Surgery of Huddinge Hospital, Karolinska Institute, Huddinge, Sweden.

Requests for reprints should be sent to Dr. Bo Bergman, Department of Psychiatry, Karolinska Institute/Huddinge Hospital, S-141 86 Huddinge, Sweden.

This paper was submitted to the journal May 5, 1990, and accepted with revisions April 4, 1991.

mented with data from medical records. The battered women and the control subjects were further compared concerning sociodemographic data collected from the population register. At the time of the follow-up 5 years after the battering, one woman had died (not from abuse, however) and another five had requested confidentiality concerning marital status, address, and nationality. Thus, the study group was reduced to 111 regarding sociodemographic data.

Statistics

For intergroup comparisons, the chi-square or Fisher's Exact Test were used.

Results

Sociodemographic Data

The marital status of the battered women was characterized by a great proportion of divorcées and a small proportion of married women compared with the control subjects (Table 1). These conditions were stable in time. However, there were significantly more changes in marital status among the battered women than among the control group. Table 2 illustrates this higher turnover among the battered women, who went through more divorces than the control subjects both during the 5 years preceding the battering and during the following 5 years. The battered women were also less prone to marry during the interval of time studied.

Somatic Hospital Care

The battered women had a significantly higher consumption of somatic hospital care during the 15-year study period (Table 3). Ninety of 117 battered women (77%) had been admitted to the hospital at least once, compared with 58 of the 117 control subjects (50%). Even more striking was the difference in the number of admissions—420 for the battered women and 119 for the control group. The most pronounced differences were found in the number of admissions due to suicide attempts, for observation (undefined disorders), and due to trauma.

Figure 1 illustrates the considerably higher use of somatic hospital care by the battered women than by the control group in all age ranges except the youngest group. The difference is most pronounced between 30 and 40 and then gradually diminishes.

TABLE 1—Marital Status of Battered Women and Control Subjects at the Time of the Battering and 5 Years Later

	Battered Women		Controls	
Marital Status	1983–1984 (n = 117)	1988–1989 (n = 111)[a]	1983–1984 (n = 117)	1988–1989 (n = 117)
Married	23*	20*	49	57
Divorced	47*	50*	17	19
Never married	44	38	48	38
Widowed	3	2	3	3

Note. *P* values denote differences between battered women and control subjects.
[a]Five battered women had confidential data and one had died.
*$P < .001$.

TABLE 2—Changes in Marital Status of Battered Women and Control Subjects 5 Years Before to 5 Years After the Battering

	Battered Women		Controls	
Changes in Status	1978–1983 (n = 117)	1984–1989 (n = 111)[a]	1978–1983 (n = 117)	1984–1989 (n = 117)
No. of divorces	16**	19**	7	9
No. of marriages	5*	13	13	14

Note. *P* values denote differences between battered women and control subjects.
[a]Five battered women had confidential data and one had died.
*$P < .05$.
**$P < .01$.

TABLE 3—Use of Somatic Hospital Care during a Period of 15 Years (1973–1988) by Battered Women and Age-Matched Control Subjects

	Battered Women (n = 117)		Controls (n = 117)	
Reason for Care	No. of Women	No. of Admissions	No. of Women	No. of Admissions
Surgical disorders (not trauma)	25	40*	19	22
Trauma	47***	70***	15	18
Gynecological disorders	48***	91***	24	41
Induced abortion	20**	22**	7	8
Medical disorders	40***	71***	14	17
Suicide attempts	23***	55***	2	2
Observation	28***	71***	9	11
Total	90***	420***	58	119

*$P < .05$.
**$P < .01$.
***$P < .001$.

Psychiatric Care

The total use of psychiatric care by the battered women during the 15-year period widely exceeded that of the control subjects. A total of 69 battered women had been admitted to a psychiatric clinic during the period in question; only 1 control woman had been admitted, and only on one occasion. Fifty-six battered women had attended a psychiatric outpatient clinic compared with 9 control subjects ($P < .001$).

In Table 4 the spectrum of psychiatric diagnoses concerning inpatients is presented. Substance abuse, suicide attempts, and depression were common. Psychoses were the least frequent major psychiatric diagnostic category.

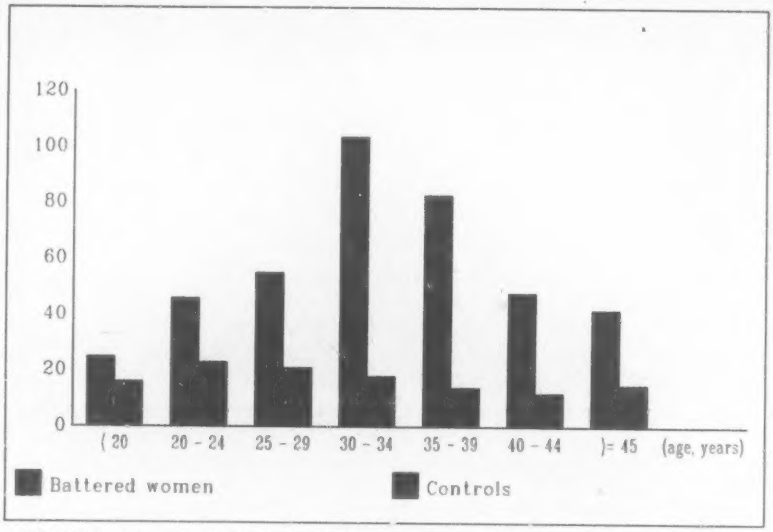

FIGURE 1—Number of admissions in somatic hospital care by battered women and control subjects in different age groups during 15 years (1973–1988).

TABLE 4—Diagnosis in Psychiatric Inpatient Care by Battered W[illegible]d Control Subjects during 15 years (1973–1988)

Diagnoses	Battered Women (n = 117) No.	Battered Women (n = 117) %	Control Subjects (n = 117)
Depression	19	16**	0
Psychoses	10	9*	1
Alcoholism	27	23**	0
Drug addiction	10	9*	0
Suicide attempt	18	15**	0
Other and unspecified diagnosis	13	11*	0
Total no. of psychiatric inpatients	69	59	1

*$P < .01$.
**$P < .001$.

Treated versus Untreated Group

Half of the 117 battered women entered the treatment program. However, when the two 5-year periods before and after the battering were compared, those who entered the program did not reduce their use of either somatic or psychiatric care.

Discussion

The present study has elucidated the battered woman's exposed situation. The women included in the study were identified in the emergency room when seeking aid for acute injuries, which presented an opportunity for early identification and intervention. However, there could be a question as to whether this way of identifying the subjects produces a study group that is only partially representative of all battered wives. The general impression made by the present study group is that it is more heavily loaded psychosocially with personal problems, including alcohol and drug abuse, than most populations of battered wives described in the literature.[4–6]

There were more divorcées among the battered women than among the control group, and during the 5 years following the battering in question the battered women continued to change their marital status more frequently than the control subjects. These facts reflect the ongoing strife in their marriages. Women who frequently change male partners probably increase their risk of being battered.

The use of somatic hospital care was dramatically higher among the battered women than among the control women. The most impressive difference between the groups concerns repeated periods of care, especially for undefined disorders, suicide attempts, and trauma. These findings are compatible with the description of the "battered wife syndrome" comprising surgical, psychosomatic, and psychiatric symptoms such as depression, headache, insomnia, and anxiety.[7–11] Although trauma is one major diagnostic category in the present study, it is not the most frequent one. This indicates that injury is just one part of a symptom complex that battered wives show. The high use of medical care could be presumed to reflect the degree of ongoing battering in the home quite well. This hypothesis is supported by the results presented in Figure 1, with a peak in use of care in the age range 25 through 39 years. Many studies have found a median age of battered wives of 30 to 35 years,[12–14] with the husband a couple of years older. The risk of being battered seems to be most pronounced in the late 20s and in the 30s.

There was no decrease in use of somatic hospital care by the women who entered the treatment program during the 5-year follow-up period. The use of somatic hospital care was stable at a high level in the entire battered group compared with the controls. The use of psychiatric care also seemed stable at a much higher level than expected. ☐

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Mammography in a Health Maintenance Organization

John P. Hansen, MD, MSPH, Penny A. Knapp, RN, BSN, and Polly A. Newcomb, PhD

ABSTRACT

The charts of a random sample of 496 women, aged 51 and older, in a staff model health maintenance organization (HMO) were retrospectively audited for mammography frequency over an 8-year period. Of the 496 women, 306 (61.7%) had at least one mammogram. For the 422 women who had been members of the HMO for at least 2 years, 196 (46.4%) had mammography within the previous 18 months. (*Am J Public Health.* 1991;81:1489–1490)

Introduction

There is little documentation in the medical literature on the performance level of health maintenance organizations (HMO) in carrying out screening mammography according to recommended guidelines. The purpose of this article is to report the frequency of screening mammography in a staff model HMO.

Methods

A 50% random sample of the female members of a staff model HMO in Madison, Wis (total membership = approximately 26 000), who were at least 51 years old and members of the HMO for at least 1 year as of March 1, 1989, was identified. Women with breast cancer diagnosed prior to any screening in this HMO or diagnosed prior to March 1, 1981, were excluded from the sample. A chart review was conducted on this sample's mammography utilization for the 8-year period from March 1, 1981, through February 28, 1989. The following information was abstracted from the chart: number and dates of mammograms, abnormal findings (defined as those resulting in a recommendation other than routine follow-up in one year), and follow-up studies. Any mammogram obtained within 6 months of a previous mammogram was excluded because it was assumed that it would have been performed as a recommended follow-up exam and not as a routine screening mammogram.

For each member, an expected number of screening mammograms was calculated, based on the member's age or length of membership. The expected number based on age was calculated by subtracting 50 from the age in years as of March 1, 1989. The expected number based on membership duration was the actual length of membership in years. The smaller of the two calculations was the expected number of screening mammograms for the member, with the maximum number expected set at 7.

No special programs, such as patient reminders, were in place to encourage mammography. There were no out-of-pocket costs to the patients for mammography.

Results

A total of 1030 women aged 51 years and older were enrolled in the HMO at the time of the study. An approximate 50% sample of 511 was randomly selected. Fifteen of these women had a previous diagnosis of breast cancer and were excluded; the final sample numbered 496. The mean age of the women was 59.4 years (range = 51–84) and the mean length of membership was 4.7 years (range = 1–13) (Table 1).

Of the 496 women, 306 (61.7%) had at least one mammogram and 78 (15.7%) had all of their expected number of mammograms. During the most recent year of the study, 178 (35.9%) had at least one mammogram. Of the 422 women who had been members for at least 2 years, 196 (46.4%) had a mammogram within the previous 18 months.

Older women were more likely to have had a mammographic examination than were younger women (P = .034). In addition, increasing length of membership

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John P. Hansen and Penny A. Knapp are with Group Health Cooperative, Madison, Wis. Polly A. Newcomb is with the University of Wisconsin Clinical Cancer Center, Madison, Wis.

Requests for reprints should be sent to John P. Hansen, MD, MSPH, Medical Director, Group Health Cooperative, One South Park Street, Madison, WI 53715.

This paper was submitted to the journal June 27, 1990, and accepted with revisions April 2, 1991.

TABLE 1—Women with at Least One Mammogram by Age and Length of HMO Membership

	Members <4 yrs		Members ≥4 yrs		Total	
Age	n	Mamm. (%)	n	Mamm. (%)	n	Mamm. (%)
51–55	68	28 (41.2)	88	60 (68.2)	156	88 (56.4)
56–60	53	22 (41.5)	88	66 (75.0)	141	88 (62.4)
61–65	51	25 (49.0)	67	50 (74.6)	118	75 (63.6)
≥66	26	14 (53.8)	55	41 (74.5)	81	55 (67.9)
Total	198	89 (44.9)	298	217 (72.8)[a]	496	306 (61.7)[b]

Note. Mamm. = number of women with at least one mammogram.

[a] χ^2 for the difference in proportion by membership length = 39.1 (1df), $P < .001$.

[b] χ^2 for the difference in proportion by age = 3.37 (3df), $P = .034$.

was strongly associated with mammography ($P < .001$) (Table 1). Of those who were between the ages of 66 and 70 and had been members in the HMO for 4 or more years, 84.6% had at least one mammogram, whereas only 41.2% of those aged 51 to 55 who were members for less than 4 years had mammography during their membership.

Discussion

The intent of this study was to document the frequency of screening mammography in women over age 50 in a staff model HMO. Three recent studies in clinical practices have reported frequency rates of women having mammography in the year prior to the study of 15.6% to 26.6%, in contrast to the 39.5% found in this study.[1–3] We also looked at the frequency of mammography within the past 18 months for women who had been members of the HMO for at least 2 years and found 46.4% of these women had received mammography. In several population surveys, frequency of mammography during the year previous to the survey ranged from 17% to 46%.[4–9] In several studies and surveys reporting the percent of women over age 50 who had ever had at least one mammogram, the frequency ranged from 39.2% to 74.0%, as compared to 61.7% in this study.[1,6–12]

It is somewhat difficult to compare the results in this study with all of the above mentioned surveys because they were conducted in different years and it is clear that the frequency of mammography has been increasing in the last several years.[4,5,12] However, given the time frame of this study, the use of mammography by this HMO compares quite favorably with that reported in the above studies.

In one clinical study,[2] it was noted that 40% of the patients who had HMO insurance had mammography in the previous year in contrast to only 16% of all patients in the study. Several studies have demonstrated an increased use of preventive services and visits by patients enrolled in HMOs compared to other health care systems.[13–16] Whether this is related to the characteristics of the HMO itself or to the reduced financial barrier characteristic of HMOs is unclear, but several other studies have shown that use of preventive services increases as patient out-of-pocket costs decrease.[8,17–19] However, one of the above studies was a well-controlled prospective study that looked at both factors and found that patients in a prepaid system had significantly higher numbers of preventive visits than those in a fee-for-service system, even when there were no out-of-pocket costs for those in the fee-for-service system.[13] It is likely that both the absence of financial barriers and the nature of the managed care system contributed to the observed frequency of mammography utilization in this staff model HMO. ☐

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Geographic Variation in Relative Fees under Medicare

José J. Escarce, MD, AM

ABSTRACT

Little is known about geographic variation in Medicare's relative fee structure. Using 1986 Part B Medicare claims data, ratios among physician fees for surgical procedures belonging to small families of closely related procedures, excluding outliers, were found to vary up to twofold. Under Medicare's current system of physician payment, physicians in different areas face different financial incentives for performing one procedure in preference to possible alternatives. Changes in incentives under a resource-based Medicare fee schedule will be more pervasive than previously recognized. (*Am J Public Health*. 1991;81:1491–1493)

Introduction

Substantial geographic variation in Medicare physician fees has been documented.[1–5] In contrast, little is known about geographic differences in the relationships among fees for different services (i.e., in Medicare's relative fee structure), although recent findings suggest that these differences may be considerable. Researchers have found that fees for different services are not highly correlated across areas,[3] and that areas with high fees for one service often have low fees for other services.[4,5]

Under the Omnibus Budget Reconciliation Act of 1989, a resource-based Medicare fee schedule will be phased in over 5 years beginning in 1992.[6] The fee schedule will impose a uniform relative fee structure in all geographic areas. Therefore, studying existing geographic differences in relative fees can shed light on the extent to which the financial incentives facing physicians will change under the fee schedule. This article describes the geographic differences in the relationships among Medicare fees for surgical procedures belonging to small families of closely related procedures.

Methods

The data for this study were obtained from the 1986 Part B Medicare Annual Data (BMAD) Procedure File. This file contains information on all physician services billed to Medicare, including each service's frequency and total allowed Medicare charges, reported at the level of Medicare charge locality and physician specialty. Services are identified using Physicians' Current Procedural Terminology (CPT) codes.[7]

Seven families of procedures were selected for analysis based on high frequency of use and representation of the principal specialties that provide services to Medicare enrollees (Table 1). Relative fees for the procedures in a family were calculated in three steps. First, each procedure's average allowed charge was determined by charge locality and physician specialty. Claims for assistance at surgery were excluded from these calculations. Second, a "reference" procedure was chosen in the family. (The choice of reference procedure is arbitrary and does not affect the study's findings.) Third, the relative fee for each procedure in each locality was calculated as the ratio of the average allowed charge for the procedure to the average allowed charge for the reference procedure in the family.

All analyses of geographic variation in relative fees were specialty specific. To reduce the influence of infrequent events on the results, the analyses for each family used only data for charge localities where each procedure in the family was performed at least 10 times by physicians in the relevant specialty. (Limiting the analyses to charge localities where each procedure was performed at least 25 times does not affect the findings.) Data for Alaska, Hawaii, and Puerto Rico were excluded.

Results

Relationships among fees for procedures in a family varied considerably across charge localities for all families studied. For the five families of procedures performed predominantly by physicians in one specialty (Table 2), relative fees varied by a factor of 1.25 to 2 even excluding the 5% of localities with the highest and lowest relative fees. Relative fees for knee arthroplasty exhibited the greatest variation. The ratio of the average allowed charge for total knee replacement to the average allowed charge for single compartment arthroplasty was less than 0.89 in 5% of localities where orthopedic surgeons performed each procedure at least 10 times, but it exceeded 1.83 in another 5% of localities. Relative fees for

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The author is with the Section of General Internal Medicine, Department of Medicine, School of Medicine, and the Leonard Davis Institute of Health Economics, University of Pennsylvania, Philadelphia.

Requests for reprints should be sent to José J. Escarce, MD, 3615 Chestnut Street, Suite 314, Philadelphia, PA 19104-2676.

This paper was submitted to the journal July 24, 1990, and accepted with revisions April 24, 1991.

TABLE 1—Families of Surgical Procedures Selected for Study

Family and Type of Procedure[7]	CPT Code[7]	Frequency[a]	Average Allowed Charge[b]
Mastectomy, with axillary lymphadenectomy			
Partial	19 162	2729	$ 851
Radical, including pectoral muscles	19 200	3231	1004
Modified radical, leaving pectoral muscles	19 240	40 314	963
Knee arthroplasty, condyle and plateau			
Medial or lateral compartment	27 446	2968	1775
Both compartments (total knee replacement)	27 447	50 882	2236
Permanent pacemaker insertion, transvenous[c]			
Atrial	33 206	17 656	979
Ventricular	33 207	35 770	996
Atrioventricular sequential	33 208	14 370	1259
Coronary artery bypass, autogenous graft			
One graft	33 510	8562	2504
Two grafts	33 511	14 762	3413
Three grafts	33 512	31 046	3751
Four grafts	33 513	24 932	4019
Five grafts	33 514	10 164	4214
Six or more grafts	33 516	3034	4281
Colonoscopy, beyond splenic flexure			
Diagnostic	45 378	254 900	390
For biopsy or collection of specimen	45 380	100 706	437
For removal of polypoid lesion	45 385	145 797	599
Prostatectomy			
Transurethral resection, complete	52 601	210 999	1082
Suprapubic, subtotal	55 821	6454	1161
Retropubic, subtotal	55 831	3076	1205
Cataract extraction, with intraocular lens			
Intracapsular	66 983	139 736	1707
Extracapsular	66 984	615 079	1665

Note. Data from 1986 Part B Medicare Annual Data Procedure File.

[a] Total number of times procedure was provided to Medicare beneficiaries in study states in 1986; claims for assistance at surgery excluded.

[b] Total allowed Medicare charges in study states divided by total frequency.

[c] Data for four localities in southeastern New York State excluded due to use of local procedure codes rather than CPT codes.

TABLE 2—Geographic Variation in Relative Fees for Five Families of Surgical Procedures Performed Predominantly by Physicians in One Specialty

Family (No. of Localities)[a] and Type of Procedure	Relative Fee (Percentile)[b] 5th	50th	95th
Mastectomy (51)[c]			
Partial	—	—	—
Radical	0.81	1.34	1.67
Modified radical	0.76	1.24	1.54
Knee arthroplasty (70)[d]			
Medial or lateral compartment	—	—	—
Total knee replacement	0.89	1.18	1.83
Coronary bypass (52)[e]			
One graft	—	—	—
Two grafts	1.02	1.34	1.76
Three grafts	1.12	1.58	1.99
Four grafts	1.17	1.70	2.23
Five grafts	1.22	1.73	2.41
Six or more grafts	1.15	1.79	2.48
Prostatectomy (61)[f]			
Transurethral resection	—	—	—
Suprapubic	0.89	1.00	1.15
Retropubic	0.81	1.01	1.14
Cataract extraction (190)[g]			
Intracapsular	—	—	—
Extracapsular	0.94	1.02	1.15

Note. Data from 1986 Part B Medicare Annual Data Procedure File.

[a] Number of localities where each procedure in the family was performed at least 10 times by physicians in the relevant specialty.

[b] Relative fee in 5% of charge localities is less than 5th percentile, in 50% of localities less than 50th percentile, in 95% of localities less than 95th percentile.

[c] Reference procedure for calculating relative fees = partial mastectomy; specialty = general surgery.

[d] Reference procedure for calculating relative fees = single compartment arthroplasty; specialty = orthopedic surgery.

[e] Reference procedure for calculating relative fees = coronary bypass with one graft; specialty = thoracic surgery.

[f] Reference procedure for calculating relative fees = transurethral resection; specialty = urology.

[g] Reference procedure for calculating relative fees = intracapsular extraction; specialty = ophthalmology.

cataract extraction with lens insertion showed the smallest variation.

Findings were similar for the two families of procedures performed in appreciable numbers by physicians in more than one specialty. Relative fees for permanent pacemaker insertion varied by a factor of 1.5 to 2 (Table 3). Relative fees for colonoscopy varied by a factor of 1.25 to 1.67 (Table 4).

Discussion

This study documents considerable geographic differences in relationships among Medicare fees for surgical procedures even when the procedures are closely related. These results extend previous findings of large differences in the ratios of the fees for several common but unrelated physician services to the fee for inguinal hernia repair.[8]

The causes of geographic variations in relative fees are unknown. One possibility is that these variations are due to market factors unique to each procedure. However, it seems unlikely that market factors could differ enough among procedures in the same family to account for the large variations found in this study. Another possibility is that the variations are caused by physician pricing decisions that are largely arbitrary and cannot be explained by the usual market forces. Further research is needed in this area.

The findings of this study suggest that physicians in different areas face different financial incentives for providing one procedure in a family in preference to others. Whether these incentives affect physician decision making is unknown. They have potential clinical relevance, however, because for many patients seen in clinical practice, the best choice among procedures in a family may not be clearcut (e.g., partial mastectomy vs modified radical mastectomy) and could be influenced by nonclinical factors. Determining if there is a relationship between geographic differences in relative fees and variations in the use of physician services by Medicare enrollees is an important subject for future research.[9,10]

This study also reveals the extent to which the financial incentives facing phy-

TABLE 3—Geographic Variation in Relative Fees for Permanent Pacemaker Insertion, by Specialty

Specialty (No. of Localities)[a] and Type of Pacemaker	Relative Fee (Percentile)[b] 5th	50th	95th
Thoracic surgery (69)[c]			
Atrial	—	—	—
Ventricular	0.80	1.01	1.20
Atrioventricular sequential	0.89	1.21	1.48
Cardiovascular medicine (67)[c]			
Atrial	—	—	—
Ventricular	0.82	1.04	1.39
Atrioventricular sequential	0.98	1.29	1.86
General Surgery (55)[c]			
Atrial	—	—	—
Ventricular	0.72	1.04	1.28
Atrioventricular sequential	0.86	1.28	1.71

Note. Data from 1986 Part B Medicare Annual Data Procedure File.

[a] Number of localities where each procedure in the family was performed at least 10 times by physicians in the specialty.

[b] Relative fee in 5% of charge localities is less than 5th percentile, in 50% of localities less than 50th percentile, in 95% of localities less than 95th percentile.

[c] Reference procedure for calculating relative fees = atrial pacemaker.

TABLE 4—Geographic Variation in Relative Fees for Colonoscopy, by Specialty

Specialty (No. of Localities)[a] and Type of Colonoscopy	Relative Fee (Percentile)[b] 5th	50th	95th
Gastroenterology (158)[c]			
Diagnostic	—	—	—
For biopsy or collection of specimen	0.99	1.12	1.26
For removal of polypoid lesion	1.18	1.48	1.90
Internal medicine (162)[c]			
Diagnostic	—	—	—
For biopsy or collection of specimen	0.96	1.11	1.35
For removal of polypoid lesion	1.14	1.47	1.92
General Surgery (149)[c]			
Diagnostic	—	—	—
For biopsy or collection of specimen	0.95	1.15	1.52
For removal of polypoid lesion	1.18	1.52	1.97

Note. Data from 1986 Part B Medicare Annual Data Procedure File.

[a] Number of localities where each procedure in the family was performed at least 10 times by physicians in the specialty.

[b] Relative fee in 5% of charge localities is less than 5th percentile, in 50% of localities less than 50th percentile, in 95% of localities less than 95th percentile.

[c] Reference procedure for calculating relative fees = diagnostic colonoscopy.

sicians will change under a resource-based Medicare fee schedule. It is widely recognized that the fee schedule will result in higher fees for evaluation and management services and lower fees for most surgical procedures.[6] What has gone unnoticed, however, is that by imposing a uniform relative fee structure in all payment areas, the fee schedule also will lead to innumerable changes in the relationships among fees for different surgical procedures. Some of these changes may be substantial.

The fee schedule, therefore, will not just realign physicians' financial incentives for choosing between evaluation and management services and surgery. In many instances, it may also considerably alter physicians' incentives for choosing among surgical procedures. Evaluating whether these changes in incentives affect patterns of surgical care, including access and appropriateness, should be part of any plan to monitor the impact of Medicare physician payment reform. ☐

Acknowledgments

This study was supported by the Physician Payment Review Commission and by a Hartford Foundation Fellowship for Research on Aging. The opinions expressed are the author's and do not necessarily reflect the views of the Physician Payment Review Commission or the Hartford Foundation.

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Age- and Sex-Specific Cycles in United States Suicides, 1973 to 1985

Richard McCleary, PhD, Kenneth S. Y. Chew, PhD, James J. Hellsten, MS, and Marilyn Flynn-Bransford

ABSTRACT

Total US suicides peak on the first day of each week, in the first week of each month, and in the late spring of each year. Although these cycles are assumed to characterize all US suicides, analyses by age and sex show that the cycles occur only in a few subpopulations. Day-of-the-week effects are found almost exclusively in middle-aged suicides; day-of-the-month effects are found almost exclusively in elderly male suicides; and month-of-the-year effects are found almost exclusively in teenaged and elderly suicides. (*Am J Public Health.* 1991;81:1494–1497)

Introduction

Analyzing 129 621 US suicides recorded during 1972 to 1978, MacMahon[1] found that (1) within each year, suicides peak in late spring and drop in winter; (2) within each month, suicides peak on the fifth day and drop until month's end; and (3) within each week, suicides peak on Monday and drop to weekend lows. Although all three cycles were reported as early as the 19th century,[2,3] MacMahon confirmed their existence in a reliable data set. Even with a 7-year series (1972 to 1978, 2557 days), unfortunately, analyses of aggregated suicides cannot answer questions of heterogeneity. Analyses of a longer 13-year series (1973 to 1985, 4748 days) confirm MacMahon's basic finding of annual, monthly, and weekly cycles. But more significantly, subpopulation analyses reveal that the cycles vary by age and sex.

Data and Method

To avoid the temporal and geographic limitations of earlier studies,[4,5] individual suicide data for 1973 to 1985 were obtained from a database compiled by the National Center for Health Statistics.[6] Data from 1972 were excluded because they were a 50 percent sample. Otherwise, our data are identical to those analyzed by MacMahon.

Although the magnitude of a time series cycle can be expressed in many ways, the simplest index is the PRE (proportionate reduction in error) statistic for the periodic means. To define this index, let $\bar{S}$ be the mean number of suicides for all (4748) days and let $\bar{w}_i$, $\bar{d}_j$, and $\bar{m}_k$ be the mean number of suicides on the i^{th} weekday, j^{th} date, and k^{th} month. Then for weekdays, dates, and months

$$\text{PRE} = 1 - \sum_{t=1}^{4748} \sum_{i=1}^{7} (S_t - \bar{\omega}_i)^2/(S_t - \bar{S})^2,$$

$$i = \text{Sunday to Saturday}$$

$$\text{PRE} = 1 - \sum_{t=1}^{4748} \sum_{j=1}^{31} (S_t - \bar{d}_j)^2/ (S_t - \bar{S})^2,$$

$$j = \text{1st to 31st}$$

$$\text{PRE} = 1 - \sum_{t=1}^{4748} \sum_{k=1}^{12} (S_t - \bar{m}_k)^2/ (S_t - \bar{S}),^2$$

$$k = \text{January to December}$$

PRE varies between 0 and 1 and is interpreted as the proportion of variance in daily suicides attributable to weekdays, dates, or months.

Results

Table 1 breaks down US suicides by day of the week, date, and month. This longer series confirms MacMahon's finding of weekly, monthly, and annual cycles. We proceed first to disaggregate total periodicity by age and sex and then to disaggregate each of the three cycles—weekly, monthly, and annual—by age and sex.

Table 2 reports PREs for day of the week, date, and month estimated for 15 age groups. Under a null hypothesis, PREs follow an *F*-distribution with degrees of freedom equal to the series length (n_1 = 4748) and the number of periods (n_2 = 7 days of the week, 31 dates, or 12 months). Null hypothesis probabilities are listed beside each PRE in parentheses. With .01 significance, many of the PREs are due to chance. The PREs are graphed

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The authors are with the Program in Social Ecology, University of California, Irvine.

Requests for reprints should be sent to Richard McCleary, PhD, Program in Social Ecology, University of California, Irvine, CA 92717.

This paper was submitted to the journal August 3, 1990, and accepted with revisions March 5, 1991.

TABLE 1—US Suicides by Day of the Week, Date, and Month, 1973 to 1985

Period	Mean	SD	n
Total	75.2723	10.7421	4748
Day of the week			
Sunday	73.4735	9.9666	678
Monday	81.7953	11.6142	679
Tuesday	76.9337	10.3868	679
Wednesday	75.0236	10.3170	678
Thursday	74.2345	10.2982	678
Friday	73.5487	9.9363	678
Saturday	71.8850	9.5233	678
Date			
1	75.9231	10.3852	156
2	73.9551	10.9799	156
3	74.8974	10.5128	156
4	76.4872	10.3883	156
5	79.1987	10.5857	156
6	77.3205	10.7709	156
7	77.7564	11.9986	156
8	77.3013	9.8134	156
9	77.3910	10.4063	156
10	76.4936	10.1062	156
11	76.3718	10.7884	156
12	77.3333	10.3502	156
13	77.1218	10.8022	156
14	76.0256	10.4597	156
15	76.2564	10.1866	156
16	76.6410	10.6304	156
17	74.8782	10.2821	156
18	75.1603	9.7156	156
19	75.1987	9.8939	156
20	75.8333	11.1748	156
21	74.5000	11.5197	156
22	73.1731	11.5279	156
23	73.8462	11.3900	156
24	73.2372	10.6396	156
25	74.1795	11.4303	156
26	72.4231	10.0554	156
27	73.6090	10.3066	156
28	72.7692	10.4831	156
29	72.5205	10.8906	146
30	71.9441	9.8208	143
31	71.7912	9.8956	91
Month			
January	72.2978	10.9898	403
February	74.2916	10.3836	367
March	77.7171	10.5292	403
April	78.0667	10.8642	390
May	77.3871	11.2118	406
June	76.0821	10.6220	390
July	75.9107	10.8555	403
August	76.3375	9.7845	403
September	75.6333	10.0283	390
October	75.0571	10.0775	403
November	73.5308	10.9628	390
December	70.9404	10.1514	403

Note. Based on 267 740 male and 89 653 female suicides for which complete dates of death were listed in the NCHS Mortality Detail File; dates of death were missing for 190 male and 35 female suicides.

in Figures 1a through 1d to highlight some striking patterns of significance.

Total Periodicity

Considering all three cycles jointly (Figure 1a), periodicity is more pronounced for male than female suicide. Males younger than 25 and older than 45 appear most affected. Although PREs for females rise between ages 50 and 70, the effect is more striking among males.

Day of the Week

This cycle plays no significant role in either male or female suicides until middle age (Figure 1b). For men, the effect builds dramatically from age 41 to age 55 and then, from age 56 to age 65, drops just as dramatically. For women, the effect is less pronounced but evident nonetheless.

Date of the Month

This cycle is insignificant for males and females until old age (Figure 1c). For males, the effect appears first at age 60 and increases in size until it stabilizes at age 70. For females, the effect occurs only between ages 60 and 70 and, compared to the male effect, is relatively small.

Month of the Year

For women, this cycle is insignificant (Figure 1d). For males, the cycle is limited principally to teenage and elderly suicides. Although both young and elderly male suicides are subject to annual periodicity, the two cyclical patterns are radically different. Figure 2 plots the standardized mean suicides for males under 16 and over 80 years of age. For the younger group, suicides peak in winter and fall to lows in summer. By contrast, suicide among the very old is low in the winter but rises as summer approaches.

Discussion

This investigation used an extended time series (1973 to 1985) to describe the age- and sex-specific periodicity of suicides in the United States. The analysis shows wide variation in the periodicity of suicide rates across ages and between the sexes. First, the analysis not only confirms that periodicity affects primarily male suicide, but also suggests that its effects are most pronounced for males under age 20 and over age 40. Second, concerning the weekly cycle, the analysis goes beyond existing literature in suggesting that the Monday peak holds mainly for men aged about 40 to 65. Third, it reveals that the monthly cycle observed by MacMahon[1] is primarily due to elderly males and, to a lesser extent, elderly females. Fourth and finally, the analysis not only implies that the annual cycle is primarily a male phenomenon, but reveals that male involvement is bimodally distributed by age: young and old but no middle-aged men are affected.

Both bioclimatic and socioeconomic hypotheses have been used to explain suicide cycles.[7–12] Whether bioclimatic or socioeconomic, most hypotheses up to now have assumed that every act of suicide is subject to the same temporal forces. Our analysis shows that this assumption of population homogeneity is not warranted. On the contrary, we find that for most individuals, the risk of suicide is distributed randomly over the week, over the month, and throughout the year. Only for some individuals—or perhaps, for all individuals at particular stages of the life cycle—is the risk of suicide distributed nonrandomly. Nonetheless, the complex distribution of periodicity suggests that age- and gender-linked contingencies are involved in suicide and that gender and life cycle should be considered in attempting to understand this phenomenon. ☐

TABLE 2—Suicide Periodicity by Age and Sex, US, 1973 to 1985

	Male		Female	
Age	PRE	Prob <	PRE	Prob <
		Total[a]		
5–15	3.709	(.0000)	1.797	(.0005)
16–20	2.702	(.0000)	1.326	(.0589)
21–25	1.478	(.0154)	1.749	(.0009)
26–30	1.597	(.0046)	1.287	(.0803)
31–35	1.022	(.4127)	1.009	(.4373)
36–40	1.305	(.0695)	1.574	(.0059)
41–45	2.456	(.0000)	1.674	(.0020)
46–50	3.423	(.0000)	1.687	(.0018)
51–55	5.583	(.0000)	2.476	(.0000)
56–60	4.266	(.0000)	2.011	(.0000)
61–65	3.114	(.0000)	2.861	(.0000)
66–70	4.723	(.0000)	1.651	(.0026)
71–75	3.950	(.0000)	1.893	(.0002)
76–80	4.121	(.0000)	1.482	(.0148)
≥ 81	3.133	(.0000)	1.320	(.0618)
		Day of the Week		
5–15	.923	(.0000)	.212	(.1220)
16–20	.701	(.0000)	.180	(.2000)
21–25	.575	(.0001)	.456	(.0014)
26–30	.571	(.0001)	.185	(.1871)
31–35	.342	(.0125)	.299	(.0275)
36–40	.532	(.0003)	.259	(.0558)
41–45	1.008	(.0000)	.486	(.0008)
46–50	2.171	(.0000)	.602	(.0001)
51–55	4.061	(.0000)	.994	(.0000)
56–60	2.788	(.0000)	.630	(.0000)
61–65	1.084	(.0000)	.919	(.0000)
66–70	1.079	(.0000)	.318	(.0193)
71–75	.389	(.0051)	.186	(.1826)
76–80	.660	(.0000)	.214	(.1172)
≥ 81	.240	(.0766)	.252	(.0826)
		Date		
5–15	.707	(.2992)	.730	(.2544)
16–20	.783	(.1719)	.767	(.1947)
21–25	.728	(.2581)	.906	(.0581)
26–30	.707	(.2994)	.833	(.1137)
31–35	.447	(.8822)	.478	(.8284)
36–40	.417	(.9218)	.882	(.0725)
41–45	.840	(.1066)	.604	(.5358)
46–50	.608	(.5255)	.501	(.7823)
51–55	.912	(.0544)	.742	(.2339)
56–60	.719	(.2756)	.514	(.7535)
61–65	.872	(.0799)	1.366	(.0002)
66–70	1.604	(.0000)	.639	(.4498)
71–75	2.123	(.0000)	.789	(.1632)
76–80	1.776	(.0000)	.710	(.2932)
≥ 81	1.855	(.0000)	.531	(.7161)
		Month		
5–15	2.108	(.0000)	.848	(.0000)
16–20	1.239	(.0000)	.378	(.0823)
21–25	.177	(.6753)	.387	(.0731)
26–30	.320	(.1745)	.280	(.2739)
31–35	.234	(.4345)	.231	(.4477)
36–40	.346	(.1252)	.433	(.0380)
41–45	.601	(.0027)	.584	(.0035)
46–50	.651	(.0011)	.578	(.0039)
51–55	.611	(.0022)	.732	(.0003)
56–60	.750	(.0002)	.857	(.0000)
61–65	1.132	(.0000)	.572	(.0043)
66–70	2.024	(.0000)	.690	(.0005)
71–75	1.400	(.0000)	.916	(.0000)
76–80	1.650	(.0000)	.542	(.0070)
≥ 81	.980	(.0000)	.530	(.0086)

Note. PRE = proportionate reduction in error; Prob = null hypothesis probabilities.

[a] Due to collinearity, total may be slightly more or less than the sum of day of the week, date, and month.

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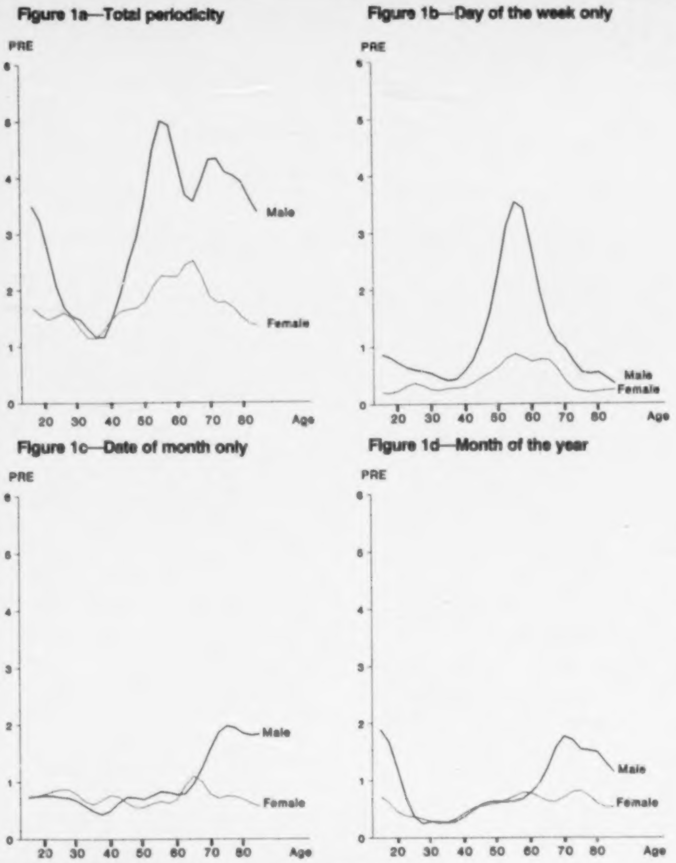

Figure 1a—Total periodicity

Figure 1b—Day of the week only

Figure 1c—Date of month only

Figure 1d—Month of the year

FIGURE 1—Proportionate reduction in error (PRE).

a—Total periodicity
b—Day of the week only
c—Date of month only
d—Month of the year

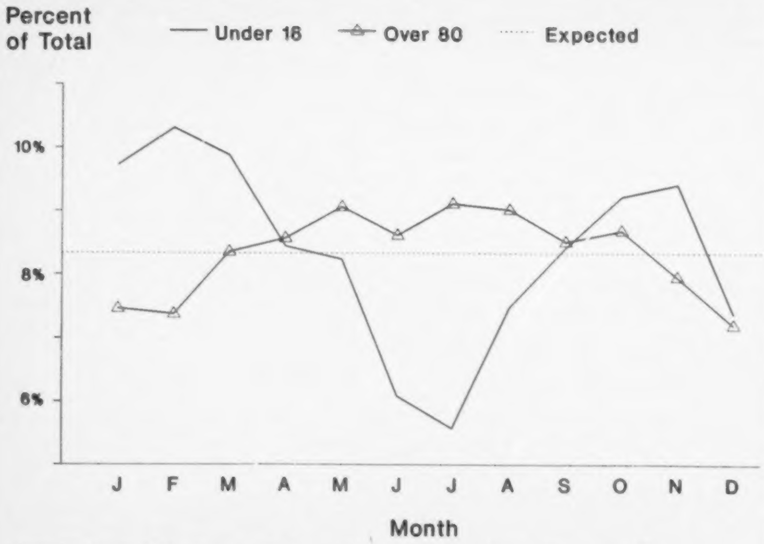

FIGURE 2—Suicides by month for males under 16 and over 80 years of age.

New from NCHS

New from NCHS presents and analyzes statistics from the National Center of Health Statistics that are relevant to public health. The department is edited by Mary Grace Kovar, DrPH. Contributions for this department should be addressed to her as follows: Mary Grace Kovar, DrPH, National Center for Health Statistics, 6525 Belcrest Road, Room 1120, Hyattsville, MD, 20782.

Family Structure and Child Health

Family structure has an important role in determining children's well-being. Accidents, injuries, and poisonings are more likely to occur in disrupted families. Children from divorced or single-parent families have more emotional, behavioral, and academic problems than children living in traditional mother–father families.

The relationship between family structure and a child's health is especially important because so many children live in nontraditional families. In 1988, 61% of all children in the United States 17 years old and younger lived with both their biological parents, 11% lived in a single-parent household with their biological mother who was formerly married to the child's father, 9% lived with a biological mother and a nonbiological father, usually a stepfather, and 8% lived in a single-parent household headed by a biological mother who was never married to the child's biological father (Figure 1). A variety of family arrangements were reported for the remaining children.

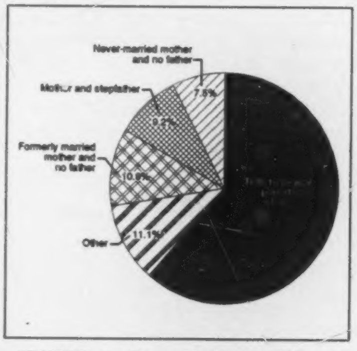

FIGURE 1—Percent distribution of children 17 years old and younger by family type, United States, 1988.

These data are from a new report based on the 1988 National Health Interview Survey on Child Health (NHIS-CH).[1] A description of the NHIS-CH is in the report. The NHIS has been described in this column before and a full description has been published.[2] In brief, the NHIS is a continuing national survey of the health of people residing in households in the United States. For the NHIS-CH, one sample child 17 years old or younger was selected from each family with a child in that age range. Information about the sample child was collected by interviewing the adult member of the family who knew most about the child's health, in most cases the child's mother. Interviews were completed for 17 110 children; the overall response rate was 91%.

Family structure was strongly associated with race and poverty status (Table 1). Children living in two-parent households, whether a nuclear family or a stepfamily, were far more likely to be White than children living in other types of families; children living with a never married mother were far more likely to be Black. Far fewer children in two-parent families lived in poverty than children in single-parent families. The proportion of children living in poverty ranged from 11% of those in a nuclear family and 14% of those living in step-families to 40% of those living with a formerly married mother, and 66% of those living with a never married mother.

Children in disrupted families were more likely to have had an accident, injury, or poisoning within the year before the interview. Children living with formerly married mothers alone or with mothers and stepfathers were more likely to have had an accident, injury, or poisoning in the past year than were other children; 17% and 16% (age-adjusted), respectively, compared with 14% of children overall.

With that exception, there was little variation in children's physical health by family type. Several conditions selected for their association with stress and other emotional problems were examined, but there were inconsistent findings for children with frequent headaches, stammer or speech defects, and bedwetting problems. Asthma was more prevalent among children living in fatherless families than in two-parent families, but the difference was not statistically significant.

However, there was a strong relationship between family type and all the indicators of emotional well-being. The indicators were school performance, rating on a behavior problems index, and treatment for emotional/behavioral problems.

Examination of school performance showed that children in nuclear families had the fewest difficulties in school. Children living with both their biological parents were about half as likely to have repeated a grade in school compared with children in all other types of families. Eleven percent (age-adjusted) of children in nuclear families repeated a grade compared with 20% of children living in stepfamilies or 21% living with a formerly married mother and 30% living with a never married mother. The same patterns were

TABLE 1—Number of Children 17 Years Old and Younger and Percent Distribution by Selected Demographic and Social Characteristics, according to Family Type, United States, 1988

Characteristic	All Children[a]	Biological Mother and Father	Formerly Married Mother and No Father	Never-married Mother and No Father	Mother and Stepfather
			Number of children in thousands		
Total[b]	63,569	38,999	6,945	4,752	5,818
			Percent distribution		
Total[b]	100.0 (0.0)	100.0 (0.0)	100.0 (0.0)	100.0 (0.0)	100.0 (0.0)
Age					
0–2 years	17.9 (0.3)	20.9 (0.4)	8.4 (0.8)	29.3 (1.3)	6.7 (0.9)
3–4 years	11.1 (0.2)	12.6 (0.4)	9.3 (0.7)	13.9 (1.1)	5.1 (0.6)
5–9 years	28.3 (0.5)	28.5 (0.6)	29.0 (1.3)	29.1 (1.4)	28.1 (1.5)
10–14 years	26.1 (0.4)	24.2 (0.5)	30.9 (1.2)	17.3 (1.2)	35.5 (1.6)
15–17 years	16.7 (0.3)	13.9 (0.4)	22.3 (1.1)	10.4 (0.8)	24.7 (1.2)
Race					
White	84.0 (0.7)	91.8 (0.5)	76.2 (1.5)	36.6 (2.0)	87.4 (1.2)
Black	16.0 (0.7)	8.2 (0.5)	23.8 (1.5)	63.4 (2.0)	12.6 (1.2)
Hispanic origin					
Hispanic	11.6 (0.5)	11.2 (0.6)	12.8 (1.0)	14.8 (1.5)	10.6 (1.0)
Non-Hispanic	88.4 (0.5)	88.8 (0.6)	87.2 (1.0)	85.2 (1.5)	89.4 (1.0)
Education of mother figure					
Less than 12 years	19.6 (0.5)	15.7 (0.6)	22.6 (1.4)	40.0 (1.8)	20.0 (1.4)
12 years	42.2 (0.5)	42.9 (0.7)	43.2 (1.5)	41.3 (1.7)	47.3 (1.6)
More than 12 years	38.2 (0.5)	41.4 (0.7)	34.2 (1.4)	18.7 (1.2)	32.7 (1.4)
Annual family income as percent of poverty line					
Less than 50%	7.6 (0.3)	3.2 (0.3)	16.1 (1.2)	34.6 (1.8)	3.1 (0.7)
50–99%	12.6 (0.4)	7.9 (0.4)	23.9 (1.2)	31.3 (1.8)	10.7 (1.0)
100–199%	27.0 (0.5)	27.1 (0.6)	27.9 (1.3)	21.7 (1.5)	32.0 (1.6)
200–299%	20.8 (0.5)	22.7 (0.5)	18.6 (1.2)	8.2 (0.8)	22.1 (1.3)
300% or more	32.1 (0.6)	39.2 (0.7)	13.4 (1.0)	4.2 (0.5)	32.1 (1.5)

Note. Standard errors of estimates are shown in parentheses.

[a] Including other and unknown family type.

[b] Including children with other and unknown values on individual demographic and social characteristics.

found for problems requiring a parent–teacher conference and for suspensions or expulsions.

Behavioral problems were assessed by a 32-item index designed to measure antisocial behavior, anxiety/depression, headstrong behavior, hyperactivity, dependency, and peer conflict/social withdrawal. The index showed that children in nuclear families had the lowest scores for behavioral problems. Correspondingly, these children were the least likely to have seen a therapist or counselor about a problem in the previous year.

These data are from 1988 when two out of five children lived in nontraditional families. However, the proportion of children who will live in such families at some time during childhood is much higher. Therefore, the risk for children is higher than it appears from a survey taken at one point in time.

The report presents findings from only one section of the 1988 NHIS-CH. Other reports from the same data set on child care arrangements; health insurance and medical care; developmental, learning and emotional problems; and children's exposure to environmental cigarette smoke have also been published.[3-6] Reports on the incidence of childhood infectious diseases, sources of medical care used by minority children, and the use of medical services by children with special health care needs are forthcoming.

The 1988 NHIS-CH is the second survey on child health included as part of the NHIS. The first child health survey was conducted in 1981. Many of the questions from the 1981 questionnaire were repeated in 1988, including items on child care arrangements; family relationships and mobility; characteristics of the child's birth, including prenatal care, hospitalization, and birthweight; chronic medical conditions; general health habits, including bedtime, sleeping arrangements, and seat belt use; school attendance and performance; and behavioral problems. New items in 1988 addressed children's potential exposure to cigarette smoke both before and after birth; developmental, learning, and emotional problems; and regular sources of medical care. ☐

Barbara Bloom, MPA
Deborah Dawson, PhD

The authors are with the National Center for Health Statistics, Centers for Disease Control, Hyattsville, Md.

Requests for reprints should be sent to Barbara Bloom, MPA, National Center for Health Statistics, Centers for Disease Control, 6525 Belcrest Road, Hyattsville, MD 20782.

Acknowledgments

The 1988 National Health Interview Survey on Child Health was developed by the National Center for Health Statistics in collaboration with the Center for Chronic Disease Prevention and Health Promotion, Centers for Disease Control; the Bureau of Maternal and Child Health and Resources Development, Health Resources and Services Administration; and the National Institute of Child Health and Human Development, National Institutes of Health.

Ordering Information

Copies of Vital and Health Statistics reports are available from the US Government Printing Office, Washington, DC 20402. Prices are given in the GPO catalogue. For Family Structure and Children's Health by D. Dawson, *Vital and Health Stat.* 10(178), order stock no. 017-022-01136-1. The price is: $2.75.

Copies of Advance Data reports are available without charge as long as supplies last from the Scientific and Technical Information Branch, National Center for Health Statistics, Centers for Disease Control, Room 1064, 6525 Belcrest Road, Hyattsville, MD 20782. Tel: 301/436-8500. Advance Data reports are also published as rapidly as possible in *Vital and Health Stat* 16 for indexing in Index Medicus and sale by GPO.

Copies of public-use computer tapes for the basic NHIS and for most data collected through the data systems of the National Center for Health Statistics are available from the National Technical Information Service, Springfield, VA 22161.

Copies of the public-use computer tapes for special studies based on the National Health Interview Survey are available from the Division of Health Interview Statistics, National Center for Health Statistics, Centers for Disease Control, Room 850, 6525 Belcrest Road, Hyattsville, MD 20782. The price for the 1981 Study of Child Health is $160. The price for the 1988 NHIS-CH is $200. Tel: 301/436-7087.

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Survey Identifies Physicians Who Manage Home Care Patients

In a recent national survey conducted by Business Mailers, Inc (BMI), a leader in healthcare information lists and databases, the company identified physicians who manage home health care patients. The study concluded that more than 31 000 doctors across all specialties oversee home health care patients.

"With so many medications being offered over the counter, consumers today are more aware of the treatments that are available to them. They are becoming more comfortable with the concept of self-care," says Mike Casey, Product Manager at BMI.

Home health care, a $1.5 billion industry, is one of the fastest growing segments in the medical field. Studies indicate that home health care is escalating at a rate of 30% per year. "This is due to such factors as the high cost of hospitalization, increasing technology, and the propensity toward self-care," says Casey.

According to BMI's study, the most commonly prescribed types of home health care are as follows: skilled nursing, unskilled nursing, home infusion therapy, and respiratory therapy.

BMI can also identify the specific types of infusion therapy (i.e., intravenous antibiotics, chemotherapy) prescribed by physicians who responded to the survey.

BMI's study also indicated that a significant amount of home health care is related to the treatment of such diseases as AIDS and cancer. It also showed the more than 17% (34 000) of the 200 000 physicians who responded treat AIDS patients, and more than 58% (111 000) treat cancer patients.

BMI also surveys doctors to determine their patient volume, prescribing patterns, over-the-counter recommendations, surgical procedures, and diagnostic test ordering frequency. The Treatment Oriented Profiles (TOP/s) database contains information on more than 240 000 physicians.

BMI's new list of doctors who manage home health care patients is available today. To obtain the list, call 1-800-888-8717 (Midwest), 1-800-524-2751 (East Coast), or 1-800-242-5478 (West Coast).